Accounting for the Gender Gap in College Attainment*

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March 5, 2010

Abstract

One striking phenomenon in the U.S. labor market is the reversal of the gender gap in college attainment. Females have outnumbered males in college attainment since 1987. We develop a discrete choice model of college entry decisions to study the driving forces of changes in college attainment by gender. We find that the increase in relative earnings between college-educated and high-school-educated individuals and the increasing parental education have important effects on the increase in college attainment for both genders but cannot explain the reversal of the gender gap. Rising divorce probabilities increase returns to college for females and decrease those for males, and thus are crucial in explaining the reversal of the gender gap in college attainment.

Keywords: gender, education, marriage, women, intergenerational schooling persistence

JEL Classification: J24, J16, I20

*We thank Michele Boldrin, Betty Daniel, Mariacristina De Nardi, Zvi Eckstein, Raquel Fernandez, John Jones, Michael Sattinger, and seminar participants at Virginia Tech, 2008 AEA meetings, University of Virginia, SUNY-Albany, 2008 North American Summer Meeting of the Econometric Society, Xiamen University, W.E. Upjohn Institute for Employment Research, the Philadelphia Fed, 2008 Annual Meeting of the Southern Economic Association, Université Laval, and 2009 Midwest Macro Meetings for helpful comments and suggestions. Ge acknowledges the AEA/CSWEP summer fellowship from W.E. Upjohn Institute for Employment Research. All remaining errors are our own.

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1 Introduction

One striking phenomenon in the U.S. labor market is the reversal of the gender gap in college attainment. In 1980, 57 percent of young men aged 25 to 34, compared with 46 percent of young women, had some college education by age 34. By 1996, however, female college attainment had reached 64 percent, 5 percentage points higher than that of males in the same cohort. In fact, females overtook males in college attainment in 1987 and have led ever since.

A large body of empirical research emphasizes the role of the earnings premium as a key explanatory variable for the determination of education outcomes (see, for example, Becker 1967; Mincer 1974; and Willis and Rosen 1979). In addition, an extensive literature shows that family background is an important determinant of the schooling decision (see, among others, Kane 1994; Cameron and Heckman 1998, 2001; Eckstein and Wolpin 1999; and Ge 2008). Recently several papers have argued empirically and theoretically that expected marriage is important in determining the schooling decision (e.g., Chiappori, Iyigun, and Weiss 2006; Iyigun and Walsh 2007; and Ge 2008).

Based on this literature, we construct a life-cycle model that includes potential costs and benefits from the labor market and marriage market which determine individual college decisions. In our model, individuals with differing learning abilities first decide whether or not to enter college. Then they might get married and have children. Parents are altruistic and value their children's learning ability, which increases with the parents' education. Forward-looking individuals take into account the impact of their own schooling on their children's learning ability. Other factors that affect an individual's decision on whether to pursue higher education include the expected direct labor market returns to college over one's lifetime, the expected marriage market returns to college, and the financial and effort costs of attending college. These costs and benefits can differ by gender.

We calculate from Panel Study of Income Dynamics (PSID) and Current Population Survey (CPS) parents' education distributions; the life-cycle profiles of single, marriage, and divorce probabilities by education; and the life-cycle profiles of earnings by education and marital status as exogenous inputs of the model. We observe that the number of college-educated parents increases over time. In the marriage market, a substantial increase of single probabilities and an increase of divorce probabilities has occurred for both genders, regardless of college attainment status. Lifetime earnings by cohort are decreasing slowly for males of all marital statuses, especially for married males. Lifetime earnings for married and divorced females are increasing gradually, while those for single females are decreasing slightly. To formally endogenize those changes is beyond the scope of our paper. We instead focus on the mechanism in which, under perfect foresight, these changes affect education decisions.

We estimate the parameters of the model by matching data on college attainment by gender from the PSID. We present evidence on how well the model fits the data. We then use the parameter estimates to simulate counterfactual experiments, which break down the sources of changes in college attainment into the effects of changes in relative earnings, changes in parental education, and changes in the marriage market.

What accounts for the increase in college attainment over the past few decades? We find that the increasing gap in earnings between college and high school graduates has important effects on the increase in college attainment for both genders. When earnings are fixed at 1946 cohort levels, attainment rates in 1996 drop by 15.5 and 14.2 percentage points for males and females, respectively. We also emphasize the importance of intergenerational persistence in schooling on the increase in college attainment for both genders. If the parents' schooling distribution is fixed at the 1946 cohort levels, college attainments in 1996 drop by 9.1 and 8.3 percentage points for males and females, respectively. The model endogenously generates the pattern that a college-educated parent is substantially more likely to have a college-educated daughter or son than is a noncollege graduate, even after controlling for the education of the other parent. This link between parents' and children's schooling provides an intergenerational propagation mechanism: as the number of college-educated parents increases, their children become more likely to attend college. Thus, the gradual transformation of parental education acts as a mechanism to propagate changes in college attainment.

What accounts for females in the last generation overtaking males in college attainment? We find that increasing divorce probabilities are crucial in explaining the relative increase in female college attainment. The rise in divorce probabilities decreases college attainment for males and increases college attainment for females. Without the observed changes in divorce probabilities, females' college attainment would have been always lower than that of males. Two factors are relevant here. First, among married persons, the returns to college education are higher for males than those for females. Second, among divorced persons, the return to college education is higher for females than for males. As divorce probabilities increase, the returns to college for divorced females become high enough to compensate for the low returns to college for married females, and thus female college enrollment exceeds that of males.

This paper contributes to an active and growing literature on gender differences in educational attainment. Several papers have studied college enrollment and graduation by gender for one cohort. Averett and Burton (1996) focus on the gender differences in college enrollment for young individuals in 1979. Rios-Rull and Sanchez-Marcos (2002) construct a model to explain why males had higher college attainment than females in the 1970s. Jacob (2002) finds that higher noncognitive skills and college premiums among women account for most of the gender gap in higher education enrollment in 1988. Those papers focus only on one cohort and thus cannot examine the trends.

Among works that study the reversal of the gender gap in higher education enrollment over time, Anderson (2002) suggests that increasing discount rates over time have a role in explaining the gender gap in college enrollment. Charles and Luoh (2003) emphasize the effect of the uncertainty of future wages on relative schooling by gender. Those papers do not consider the effects of marriage and children on college entry decisions. Chiappori, Iyigun, and Weiss (2006) show, in a theoretical framework, that women can acquire more schooling than men if the gender wage gap narrows with the level of education. One crucial assumption of their model is that the intramarital share of the marriage surplus one can extract increases with his or her education. Our results do not rely on this assumption. Goldin, Katz, and Kuziemko (2006) show that improvement in test score and high school performance, driven by the increase of expected labor market return to education, can explain most of the relative increases in women's college completion rate. They do not quantify different returns to education. To our knowledge, our paper is the first that incorporates several factors in a structural model to quantitatively account for the reversal of the gender gap in college attainment.[1]

The paper is organized as follows. In Section 2, we present some empirical results from the PSID documenting college attainment rates in 1980–1996. In Section 3, we present our model. Section 4 provides parameters estimated from the data that are used in the model. Section 5 presents the quantitative results of the benchmark model and investigates the quantitative importance of changes in relative earnings, changes in parental education, and changes in the marriage market. Brief concluding remarks are provided in Section 6.

2 Data on College Attainment

We use the PSID to calculate college attainment rates. The PSID is a longitudinal survey of U.S. families and the individuals who make up those families. We select individuals in the core sample whose ages were between 25–34 in that year and who had valid information on parents' education.[2] We use completed schooling among mature adults as the measure of an individual's schooling.[3] An individual who has more than 12 years of education completed by age 34 is defined as having a college education. The college attainment rate is calculated as the fraction of individuals that have college education among each specific group.

[1]Sanchez-Marcos (forthcoming) quantifies the reduction in gender gap in college attainment in a structure model. She does not study the overtaking of female college attainment afterwards.

[2]We thus use the average college attainment of 10 birth cohorts. The sample size in the PSID is too small for us to analyze each birth cohort.

[3]See Charles and Luoh (2003) for a discussion of the advantage of using school attainment among mature adults over enrollment.

Figure 1 illustrates the changes in relative college attainment by males and females over the sample period considered here, 1980 to 1996.[4] Among those whose ages were between 25 and 34 in 1980, 57 percent of young men had some college education, which was 11 percentage points higher than those for young women. By 1996, male college attainment rate had increased slightly by 2 percentage points, while female college attainment had increased by 18 percentage points. In fact females have led males in college attainment since 1987.[5, 6]

We also calculate college attainment rates conditional on parents' education. A detailed description of the data processing procedure is provided in Appendix 7.1. Figure 2 shows female college attainment conditional on parental education. We observe that a college-educated parent is substantially more likely to have a college-educated daughter than is a noncollege-educated parent, even after controlling for the education of the other parent.[7] For example, among those who were in the age range of 25 to 34 in 1980, 84 percent of females whose parents both had a college education had attended some college, which was 5 percentage points higher than those whose father had a college education but whose mother did not, and 20 percentage points higher than those whose mother had a college education but whose father did not. Therefore, the marginal effect of a father's education on his children's education is larger than that of a mother's. We also observe that the conditional attainment rates increase at a much slower pace than does the aggregate attainment rates. This indicates that a large fraction of the observed increase in female attainment can be accounted for by the increase in their parents' attainment.

The schooling distribution of our PSID sample's parents is shown in Figure 3. We observe

[4] We choose this beginning period to avoid the high male-to-female ratio in the early 70s after the Vietnam War. We choose this ending period because of the availability of data. The latest year of data available to us is 2005 PSID. Since we use education completed by age 34, individuals at the age of 34 in 2005 were 25 in 1996. For the years 1997 and later, of course, education by age 34 is not available for individuals at age 25.

[5] Other studies (see, for example, Charles and Luoh 2003; Goldin, Katz, and Kuziemko 2006), which use different measures of education or different data sets, find similar patterns.

[6] The sample size in PSID is too small if we divide the sample by race/ethnicity. The process of convergence and ultimate ascendancy by women in completed schooling among successive generations of men and women is evident, however, when we divide sample by race/ethnicity using the CPS.

[7] Similar patterns hold for sons, and the results are available from the authors upon request.

that the number of college-educated parents increases over time. In 1980, 12 percent of individuals ages 25–34 had parents that both had college educations, and 69 percent had parents that both had only high school educations or below. By 1996, the fractions have changed to 23 percent and 50 percent, respectively.

3 The Model

The economy is a discrete-time overlapping generations world. We assume that going to college entails an idiosyncratic nonpecuniary effort cost $D \in [0, \infty)$.[8] Adult population at age 18 is characterized by a distribution of effort costs. At age 18, individuals with different costs make schooling decisions. Each period, they might get married and have children. Parents are altruistic and care about their children's learning ability. We assume that the higher a parent's education, the higher is his or her children's learning ability. Factors that affect an individual's decision on whether to attend college include the direct labor market returns to college, the marriage market returns to college, the impact of one's own schooling on his or her children's learning ability, as well as the effort cost. These costs and benefits can differ by gender. We now describe the model in more detail.

3.1 Labor Market and Marriage

Each period, individuals of schooling type s_f and s_m might marry at an exogenously given probability, and they work.[9] Let z denote marital status, where 0 stands for being single, 1 stands for being married, 2 stands for being divorced. Let $y^{z=1}_{c,t,g,s_m,s_f}$ denote the earnings of a married individual born in year c, at age t, of gender $g = \{f, m\}$, the education

[8]We can interpret the effort cost as net of the psychic benefit of attending college. Heterogeneity in effort cost in our model is equivalent to heterogeneity in the consumption value of school in the literature (Keane and Wolpin 1997, 2001; Eckstein and Wolpin 1999; and Ge 2008). These papers consider the life-cycle decisions of one cohort. They normally allow individual heterogeneity in other dimensions, for example, different wage offers. However, Ge (2008) shows that heterogeneity in the consumption value of school is the most important determinant of women's college enrollment decision.

[9]For simplicity's sake, we do not model marriage as a match outcome. Fernandez, Guner, and Knowles (2005) study the interactions between household matching, inequality, and per capita income.

of husband is denoted by $s_m = \{1 \text{ (high school)}, 2 \text{ (college)}\}$, and the education of wife is denoted by $s_f = \{1, 2\}$. A single or divorced individual's income only depends on his/her own characteristics. For example, $y^{z=0}_{c,t,m,s_m}$ denote the earnings of a single male born in year c, at age t, of education level s_m.

We assume that fertility is exogenous. The cost of having children as the opportunity cost of time will be incorporated into our estimates of the earnings process. The financial costs of raising children is captured by household equivalence scale function $\eta(x)$ that converts household consumption into individual consumption, where x is the number of persons in the household.

The learning ability of a couple's children, a', is a function of the couple's human capital, s_m and s_f. The production function of children's learning ability is Cobb-Douglas

$$a'_{s_m,s_f} = \log[s_m^{1-\theta_s} s_f^{\theta_s}]. \tag{1}$$

This functional form captures the fact that when parents are more educated, their children tend to have high learning ability.[10] This could occur because more educated parents provide a better environment for children to flourish, or because parental learning ability is passed on genetically (Plug and Vijverberg 2003). Children of different genders from the same family have the same learning ability.

We allow for out-of-wedlock child bearing and these children live with their mother. We assume that only the mother derives utility from the ability of her out-of-wedlock children. Let $n = 2$ denote the case of having children, $n = 0$ denote the case of no children. The momentary utility function for a single female and a single male, of schooling s is, respectively,

$$u_c^{z=0,f}(t_f = t, s_m, s_f = s, n) = \frac{y^{z=0}_{c,t,f,s}}{\eta(n+1)} + n\lambda_a a'_{s_m,s_f} + \delta_z, \tag{2}$$

$$u_c^{z=0,m}(t_m = t, s_m = s, n) = y^{z=0}_{c,t,m,s} + \delta_z, \tag{3}$$

where λ_a measures the weight on the utility from children's learning ability.[11] The term

[10] In addition, for couples who are both high school graduates, the ability of their children is normalized to be 0.

[11] An alternative specification of altruism is to use a dynastic model. This model would complicate our analysis significantly because the environment is not stationary. Thus we use children's learning ability to approximate their expected lifetime utility. Learning ability is an important component of an individual's

δ_z is the value from being at state z, which can be negative or positive.

When a single mother marries, the husband only derives utility from the ability of his own children, while the wife derives utility from the average ability of all her children. Let n_{-1} denote the number of children born before marriage, a'_{-1} denote the ability of children born before marriage. The utilities of men and women at a marriage type $\left(t_m, t_f, s_m, s_f, n, n_{-1}, a'_{-1}\right)$ are given by

$$\begin{aligned} u_c^{z=1,f}(t_m, t_f, s_m, s_f, n, n_{-1}, a'_{-1}) &= \frac{y^{z=1}_{c,t_m,m,s_m,s_f} + y^{z=1}_{c,t_f,f,s_m,s_f}}{\eta(n+n_{-1}+2)} + \frac{n\lambda_a a'_{s_m,s_f} + n_{-1}\lambda_a a'_{-1}}{(n+n_{-1})/2} + \delta_z, \\ u_c^{z=1,m}(t_m, t_f, s_m, s_f, n, n_{-1}) &= \frac{y^{z=1}_{c,t_m,m,s_m,s_f} + y^{z=1}_{c,t_f,f,s_m,s_f}}{\eta(n+n_{-1}+2)} + n\lambda_a a'_{s_m,s_f} + \delta_z. \end{aligned}$$

In case of divorce, children (include children born out of wedlock) live with the mother. The mother gets a share of the husband's income, $\lambda_d \in [0,1]$. The father still derives utility from the ability of his own children. The momentary utility function for a divorced female and a divorced male, is,

$$\begin{aligned} u_c^{z=2,f}(t_m, t_f, s_m, s_f, n, n_{-1}, a'_{-1}) &= \frac{y^{z=2}_{c,t_f,f,s_f} + \lambda_d y^{z=2}_{c,t_m,m,s_m}}{\eta(n+n_{-1}+1)} + \frac{n\lambda_a a'_{s_m,s_f} + n_{-1}\lambda_a a'_{-1}}{(n+n_{-1})/2} + \delta_z, \\ u_c^{z=2,m}(t_m, s_m, s_f, n) &= (1-\lambda_d) y^{z=2}_{c,t_m,m,s_m} + n\lambda_a a'_{s_m,s_f} + \delta_z. \end{aligned}$$

3.2 The College Decision

The decision to go to college depends on the cost and the expected returns to college. A female individual born in year c chooses whether to attend college, $s_f = 1$ and $s_f = 2$, given her individual cost of schooling D, by solving

$$\max_{s_f} \left\{ U_c^f(s_f) - D \right\}, \tag{4}$$

where lifetime utility for a female is defined as:

$$\begin{aligned} U_c^f(s_f) = &\sum_{t_f=18}^{65} \beta^t \sum_{z,t_m,s_m,n,n_{-1},a'_{-1}} [u_c^{z,f}(t_m, t_f, s_m, s_f, n, n_{-1}, a'_{-1}) \\ &P_c\left(t'_m, t_f+1, z', s'_m, s'_f, n', n_{-1}, a'_{-1} | t_m, t_f, z, s_m, s_f, n, n_{-1}, a'_{-1}\right)] - \sum_{t_f=18}^{21} [C_{c+t_f} 1_{s_f=2}]. \end{aligned}$$

endowments when the college entry decision is made. As is shown in Keane and Wolpin (1997), variance in expected lifetime utility between endowment types could account for 90 percent of the total variance.

Note that β is the discount factor, $1_{s_f=2}$ is an indicator that takes the value of 1 if $s_f = 2$. C_{c+t_f} is the annual cost of attaining college in year $c + t_f$, and P_c is the probability of changing status. A male's problem is defined analogously.

An individual is indifferent as to whether he or she goes to college or not if the expected utility gain from going to college is equal to the effort cost D. We define the threshold levels as

$$D_c^{f*} \equiv U_c^f(s_f = 2) - U_c^f(s_f = 1), \tag{5}$$

$$D_c^{m*} \equiv U_c^m(s_m = 2) - U_c^m(s_m = 1). \tag{6}$$

Therefore, a female born in year c with an idiosyncratic effort cost D chooses to go to college, $s_f = 2$, if and only if $D < D_c^{f*}$, and a male chooses $s_m = 2$ if and only if $D < D_c^{m*}$.

3.3 Distribution

Each individual receives a draw of effort cost D at age 18. We assume that the individual's learning ability, a, affects the distribution of effort cost from which he or she draws. More specifically, we assume that the effort cost D is log-normally distributed with mean $\mu(a)$ and variance σ^2, where $\mu(a)$ is decreasing in the learning ability level a. Recall from Equation (1) that a is determined by parent's type, $a_{s_{m_{-1}}, s_{f_{-1}}}$, where $s_{j_{-1}}$ is parent j's schooling. In each period, a has 4 different values. Let $\psi_g^c(s_{m_{-1}} = i, s_{f_{-1}} = j)$ denote the college attainment rates of individuals of gender g, conditional on parents' education, which are calculated using the cumulative distribution function of D at D_c^{g*} as follows:

$$\psi_g^c(s_{m_{-1}} = i, s_{f_{-1}} = j) = F[D_c^{g*} | a_{i,j}]. \tag{7}$$

Notice that the fraction of individuals that go to college will depend on the parents' type, because the parents' type determines the average effort cost these individuals bear.

Let the total fraction of individuals born in year c of gender g attending college be Φ_g^c. Denote $p_{-1}^c\left(s_{m_{-1}} = i, s_{f_{-1}} = j\right)$ as the fraction of fathers and mothers with education level i and j, respectively. Thus the aggregate college attainment, Φ_g^c, is the average of the conditional attainment rates weighted by parents' education distribution:

$$\Phi_g^c = \sum_{i,j=1}^{2} \psi_g^c(s_{m_{-1}} = i, s_{f_{-1}} = j) * p_{-1}^c \left(s_{m_{-1}} = i, s_{f_{-1}} = j\right). \tag{8}$$

4 Data Inputs

We calculate parents' education distributions, marriage distributions, and earnings during the life cycle as inputs of the model. We compute the distribution of parents' education from the PSID. The results were presented in Section 2. Since the CPS cover longer periods and have a larger sample than the PSID, we use the CPS to estimate earnings and marriage distributions during the life cycle.[12] This section describes the estimation procedure and results of those inputs in detail.

4.1 Marriage Distributions

We estimate the probability that each individual will be single, married, or divorced from the March supplement of the CPS 1964–2007. We define an individual as single if he or she has never married. Individuals whose marital status is that of widowed, divorced, or separated is treated as divorced. We define an individual as having a college education if he or she completes more than 12 years of schooling, and we define an individual as having a high school education if he or she completes 12 years of schooling or less.

For each birth cohort, we first construct a pseudo-panel of people between the ages of 18 and 65. In each pseudo-panel we construct, we calculate the life-cycle profiles of fractions of individuals that are single, married, and divorced at each age, respectively. Usually not the entire life-cycle profile is observed.[13] We then use a polynomial in age and a cohort dummy to estimate the life-cycle profiles of percent single, married, and divorced for each type.

Figures 4 and 5 show life-cycle profiles of percent being single, married, and divorced by gender and education for selected cohorts. Those figures show that, over time, a substantial

[12]The PSID and the CPS show similar patterns of college attainment.

[13]For example, for a cohort born in 1970, the available CPS data only provide us with a marriage probability profile from ages 18 to 37.

increase of percent single and a significant decline in percent married for both genders and both education groups has occurred.[14] We observe that education delays marriage: Having attained the level of college implies a higher probability of being single before age 30 than if one had not. Percent divorced for college educated are lower than those for high school graduates. Percent divorced for both genders and both education groups are quite stable.

The observed aggregate stabilization in percent divorced can be driven by a combination of the decline in percent married and an increase in divorce probability. We show, in Figures 6 and 7, the cumulative marriage survival probabilities: For a married individual at age 20, what's the likelihood he/she stays married by age 45? Marriage survival probabilities for college-educated males are higher than those for male high school graduates. Both decline over time. However, marriage survival probabilities for college-educated females are lower than those for female high school graduates for earlier cohorts, but higher for later cohorts.[15]

We assume the decrease in single, marriage, and divorce probabilities to be exogenous.[16] In this paper, we focus on how future expected marriage status affects education decisions, when individuals take into account that going to college will change their future perspective on marriage.

We then calculate, conditional on being married, the probability of marrying each type of spouse. We use household and spousal identification information to match couples. Our results, not shown, confirm the well-known phenomenon that people do not marry randomly and that assortative matching exists (Becker 1973; Mare 1991; and Pencavel 1998).[17] A

[14]The marriage market is clear at each point in time by construction. Marriage divorce by gender differ for each cohort because a person may marry a spouse from another cohort.

[15]Note that our calculated marriage survival rate from age 20 disputes the claim that "half of all marriages end in divorce" for two reasons: (1) For those people who marry after age 20, the cumulative divorce rate is lower. (2) The cumulative divorce rate in America for first marriages is 41%, for second marriages is 60%, and for third marriages is 73%. What we report here is an average.

[16]Stevenson and Wolfers (2007) review the potential reasons to explain the changes in marriage and divorce probabilities. Greater access to birth control and abortion might reduce marriage (Akerlof, Yellen, and Katz 1996; Goldin and Katz 2002). Labor-saving technology might decrease the return to be gotten from specialization within a household. Increasing wage inequality might increase the time needed to search within the marriage market (Gould and Paserman 2003). In addition, the elimination of fault-based divorce and a shift from consent to unilateral divorce laws might have increased divorce probabilities.

[17]Benham (1974), Boulier, and Rosenzweig (1984), Behrman, Rosenzweig, and Taubman (1994), and Weiss (1997) point out that one's own schooling can improve spousal schooling acquired in the marriage market, but it is difficult to conclude whether this effect is due to human capital accumulation within the household

college-educated person is more likely to marry a college-educated spouse and benefit from the spouse's earnings.

4.2 Out-of-wedlock Birth Rate

We then calculate out-of-wedlock birth rate during the life cycle.[18] Figure 8 shows the cumulative probability of having children for single women: For a woman who stays single from age 18-30, what's the likelihood she has a single child by age 30? Out-of-wedlock birth rate is much higher for high school graduates than for college graduates. Both have increased dramatically over time.

4.3 Earnings

We need to estimate the expected life-cycle earnings profiles for each marriage status for an individual at the beginning of the life cycle. We do not observe wages for those who do not work, since there are none. If labor force participation is correlated with unobservable determinants of wages, a simple OLS regression is biased. To control for the selection bias, we use a two-stage procedure to estimate the wage: First we estimate equations of observed labor market participation as functions of explanatory variables along with random disturbance terms representing unobservable factors. Then we specify and estimate equations of the logarithm of wage, controlling for participation selection.

4.3.1 Estimation Procedure

We estimate a regression function for each subsample of working individuals by gender as

$$\log w_i = X_i\beta + \alpha V_i + \eta_i, \tag{9}$$

or assortative mating.

[18] We assume the total number of children a married couple has is independent of each spouse's education. Fernandez and Rogerson (1998), Rios-Rull and Sanchez-Marcos (2002), and Greenwood, Guner, and Knowles (2003) show that fertility declines slightly with income and education. Adopting the assumption that fertility declines with education should only change the results marginally.

where $\log w_i$ is the logarithm of real hourly wage and X is a vector of characteristics such as schooling and work experience. The variable V, the inverse Mills ratio, represents the selection effect of participation.

We apply the Heckman (1979) and Lee (1978) two-stage estimation methods to this model to obtain consistent estimates. First, we estimate equations linking observed labor market participation to a set of explanatory variables and a random disturbance term representing unobservable factors.[19] Second, we use these estimates to construct the inverse Mills ratios for the wage equation. Then, we run an OLS regression of log wage equations on X, using the estimated inverse Mills ratios as additional regressors, as is specified in Equation (9). Finally, we predict hourly wage for each individual using the fitted equation:

$$\log \widehat{w}_i \quad = \quad X_i \widehat{\beta}, \tag{10}$$

where $\widehat{\beta}$ is the consistent estimation of β. The estimation results, along with a full description of our methodology, are provided in Appendix 7.2.

4.3.2 life-cycle Profiles of Earnings

We use the following procedure to estimate the average life-cycle profiles of earnings from the CPS. For each birth cohort c, we first construct a pseudo-panel between ages 18 and 65. Then earnings is calculated as the product of mean predicted hourly wages (as in Equation [10]) and mean annual hours worked by that particular type.[20] We then use a polynomial in age, t, to estimate the life-cycle earnings profile for type $\Gamma = \{g, s_m, s_f, z\}$:

$$y_c^\Gamma(t) = \beta_0^\Gamma + \beta_1^\Gamma \cdot t + \beta_2^\Gamma \cdot t^2 + \beta_3^\Gamma I(cohort = c) + \varepsilon_c^\Gamma(t), \tag{11}$$

where $I(cohort = c)$ is a dummy for birth cohort c.

Figure 9 shows the estimated cohort effect of earnings at each marriage status by gender and education in 2006 dollars. Over time, earnings by cohort are decreasing slowly for males

[19]The standard procedure for ensuring identification is to have this set of variables not be identical to X. In our specification, the number of children is assumed to affect the participation decision, but not wages directly.

[20]Our measurement of lifetime earnings thus incorporates both the changes in labor supply and changes in wage. We do not disentangle those two forces in the data because in our model those two forces affect education decision through the same channel by changing earnings.

at all marital statuses, especially for married males.[21] Earnings for married females by cohort are increasing gradually, partially because of the increasing female labor supply and partially because of the increasing of wages.[22] A similar pattern is observed for divorced females. On the contrary, earnings for single females by cohort are decreasing gradually, since the increase of wages is offset by the decrease of labor supply.[23]

Figure 10 shows life-cycle profiles of earnings for each marriage status by gender and education in 2006 dollars for the 1946 birth cohort. We observe substantial earnings returns to education: High school graduates on average earn less than do college graduates, regardless of marriage status and gender. We also notice that married males on average earn more than single and divorced ones do.[24] The marriage premium is the highest for those whose spouses have college degrees. However, we do not find that married females–unlike males–earn more than do single and divorced females. The marriage premium for females is negligible among high school graduates and is in fact negative among college graduates.[25] Single females on average earn more than divorced ones do. Comparing earnings by gender, we see that single females earn an income similar to that earned by single males. Married females earn much less than do married males with a spouse of the same education. Divorced females earn less than do divorced males.

[21]Our finding is consistent with Kambourov and Manovskii (2005), who show that the life-cycle profiles of males' earnings for younger cohorts are lower than those for older cohorts.

[22]Existing theories that explain the increase in female labor participation include the following: technological innovation (Greenwood, Seshadri, and Yorukoglu 2005; Goldin and Katz 2002; and Albanesi and Olivetti 2006), falling child care costs (Attanasio, Low, and Sanchez-Marcos, forthcoming), an increase in the number of jobs that are less physically demanding (Goldin 1990), cultural acceptance of maternal employment (Fernandez, Fogli, and Olivetti 2004; Fernandez 2007; Fogli and Veldkamp 2008), and increases in women's wages (Jones, Manuelli, and McGrattan 2003). Existing theories that explain the decrease of the gender wage gap include gender differences in qualifications and discrimination (Blau and Kahn 2000), and self-selection (Mulligan and Rubinstein 2007).

[23]Our results confirm McGrattan and Rogerson (2004), who use census figures and find a decline in hours worked by single females.

[24]Korenman and Neumark (1991), among others, attribute most of the male marriage wage premium to productivity increased in marriage.

[25]Papers measuring marriage premium using wages generally find a negligible premium for females. We find negative marriage premium among female college graduates because single college females work more than do their married counterparts.

4.3.3 Cost of Attaining College

We set the annual cost of college based on estimates from the National Center for Education Statistics (NCES, Digest of Education Statistics, 2004, Table 313). Annual cost of college includes tuition, room, and board.

5 Findings

Can the model replicate the change in college attainment that occurred between 1980 and 1996? To determine this, we use the data reported in Section 4 and estimate the other model's parameters which are constant over time by matching college attainment rates obtained in the data. Then we run counterfactual simulations to study the effects of different mechanisms on college attainment by comparing college attainments from each simulation with those in the benchmark.

5.1 Benchmark

We use calculated life-cycle earnings, marriage distributions, and parent education distributions as inputs of the model. The discount factor β is set to be 0.9615 to match an interest rate of 4%. The term $\eta(x)$ is defined in the following table (Fernández-Villaverde and Krueger, 2007):

x	Family Size	1	2	3	4	5	6
$\eta(x)$	Equivalence Scales	1	1.34	1.65	1.97	2.27	2.57

We estimate the remaining 10 parameters by simulated method of moments (SMM). The solution of the college entry decision model serves as input into the estimation procedure. Specifically, a weighted average difference between sample moments and simulated moments is minimized with respect to parameters of the model. The weights are the inverses of the estimated variances of the moments. We describe the weighting procedure and the details of SMM estimation in Appendix 7.3. The moments we use are the college attainment rates conditional on parental education for both males and females between 1980 and 1996. We have 8 moments (4 parental types $\times$ 2 gender types) in each year for 17 years. In total, there are 136 moments. The parameter estimates and their asymptotic standard errors are presented in Table 1.

Parameters in the ability production function and in the effort cost distributions are identified from the levels and rank orders of the conditional attainment rates at any point of time. Ability production parameter θ_s is less than 0.5, indicating that fathers' education affects children's learning ability more than does the mothers' education. The fact that θ_s is less than 0.5 implies the following order of learning ability levels by parents' education: $a_{1,1} < a_{1,2} < a_{2,1} < a_{2,2}$. This in turn implies that the effort cost distribution parameters by parents' education $\mu_{i,j} = \mu(a_{i,j})$ have a corresponding rank order. In particular, since $\mu(a)$ is decreasing in a, we have $\mu_{1,1} > \mu_{1,2} > \mu_{2,1} > \mu_{2,2}$, which is key to be consistent with Figure 2 where the marginal effect of fathers' education on children's education is larger than that of the mothers'.

On the other hand, preference parameters affect the returns to college given each marital status. As λ_a increases, returns to college would increase if one expects to have children. This is because the additional benefit a college graduate can get from improving his/her children's learning ability depends on λ_a. The estimated utility values of being single and being divorced indicate that the status of being single brings up utility, but the status of being divorced brings down utility. If college and high school graduates have the same probabilities of being single and of being divorced, these utility values would have no impact on the returns to college, and thereafter no impact on college attainment rates. However, compared with high school graduates, college graduates are more likely to be single and less likely to be divorced. The values of δ_0 and δ_2 can therefore influence college attainment rates. In addition, the transfer parameter, λ_d, affects the gender difference in earnings return to college. Finally, as the marriage/single/divorce probabilities change over time, these preference parameters affect the relative importance of each change on the observed variations in the college attainment rates over time.

Parameters	Estimates	Asymptotic standard errors
Preference		
λ_a	13.0237	(0.0071)
δ_0	3.6116	(0.0127)
δ_2	-1.0424	(0.0223)
λ_d	0.2535	(0.0008)
Ability production		
θ_s	0.4796	(0.0003)
Effort cost distribution		
$\mu_{1,1}$	5.9859	(0.0002)
$\mu_{2,1}$	5.6675	(0.0004)
$\mu_{1,2}$	5.7699	(0.0003)
$\mu_{2,2}$	5.5183	(0.0004)
σ	0.3432	(0.0002)

Table 1: Parameters used in the benchmark model

Figure 11 compares aggregate college attainment rates from the model with those in the data. The model is able to generate the pattern that college attainments for females were lower in 1980 and higher in 1996 than those for males, as is observed in the data. In the model, female college attainment began to exceed that of males in 1988, one year later than was observed in the data.

Figure 12 compares females' college attainment rates conditional on parent's education from the model with those in the data. The model is able to generate the pattern that a college-educated parent is substantially more likely to have a college-educated daughter than is a parent who is a noncollege graduate, even after controlling for the education of the other parent.[26] In our model, parents' type determines the average effort cost these individuals bear. Thus the order of μ's, $\mu_{1,1} > \mu_{1,2} > \mu_{2,1} > \mu_{2,2}$, is critical to generate the order of school attainment by parent's type.

5.2 Counterfactual Simulations

In the benchmark economy, changes in college attainment over time are caused by the exogenous changes in parental education, life-cycle profiles of earnings by education, and

[26]Similar patterns hold for males, and the results are available from the authors upon request.

marriage distributions. To study the quantitative effects of different mechanisms on college attainment, we run counterfactual simulations. For each simulation, we keep the values of the variables that we want to focus on fixed in the 1946 cohort level, and we keep the values of other variables the same as in the benchmark model. Therefore, the comparison between each simulation and the benchmark model results will quantify the direct effects of those variables.

5.2.1 Parents' background

First, we investigate the intergenerational schooling effects. The results are shown in Figure 13. When the parents' schooling distribution is fixed at the 1946 cohort level, college attainment drops by 9.1 and 8.3 percentage points in 1996 for males and females, repectively. We notice the gender reversal of college attainment occurs in the same year as in the benchmark. Therefore, parental education is an important source of the increase in college attainment but cannot in itself account for the reversal of the gender gap.

The benchmark model captures the intergenerational persistence in schooling: When parents are more educated, their children tend to have high learning ability and are more likely to go to college. Thus the gradual transformation of parental schooling composition, as is shown in Figure 3, acts as a mechanism to propagate change in college attainment: as the number of college-educated parents increases, so does the proportion of children with high learning ability (a low value of the effort cost D), which then helps to increase the attainment rate of the children's generation. This propagation mechanism seems to affect females and males in similar magnitude, so that it had little effect on the timing of gender reversal of college attainment.

These results for intergenerational schooling effects are broadly consistent with previous research. Many studies report a significant positive relationship between parents' education and the schooling of their children for one cohort (Behrman 1997; Behrman and Rosenzweig 2002). Based on data from the National Longitudinal Survey of Youth 1979 (NLSY79), Ge (2008) estimates a sequential college choice model and shows that improvements in parental education can account for a large part of the college attendance difference between NLSY79 young women and those born almost 20 years later. To our knowledge, our paper is the first attempt to investigate the importance of intergenerational schooling effects in accounting for the trends of college attainment for both genders.

5.2.2 Earnings

To understand the effect of earnings on education, we calculate the labor market return to education. First, we calculate total lifetime earnings using the estimated life-cycle earnings profiles described in Equation (11). For a male of type $\Gamma = \{g, s_m, s_f, z\}$, we calculate total discounted life-cycle earnings at the beginning of his adult life, Y_c^{Γ}, as

$$Y_c^{\Gamma} = \sum_{t=18}^{65} (\frac{1}{1+r})^{t-18} y_c^{\Gamma}(t),$$

where r is the annual real interest rate, and $y_c^{\Gamma}(t)$ is the annual real earnings at age $t = \{18, 19, ..., 65\}$ as given by Equation (11).[27] A female's lifetime earnings are calculated analogously. An interest rate of $r = 4\%$ is used.

We then calculate the labor market return to education. For singles, we compute the differences in life-cycle earnings between college and high school for males and females. For married couples, the relevant concept of earnings is household lifetime earnings. For a married female (male), we compare the earnings of a household in which the wife (husband) has a college education, but the husband (wife) does not, with the earnings of a household in which both spouses are high school graduates.[28] For divorced couples, we compute the differences in life-cycle earnings between college and high school for males and females, adjusting for transfer from males to females.

Figure 15 presents the earnings return to college by gender and marital status. Several patterns are observed. First, the earnings return to college increases for both genders and for

[27] We assume college students cannot work, and thus we do not have earnings for those between the ages of 18 and 21.

[28] We also compared earnings in households where both spouses are college graduates with earnings in households in which the wife (husband) has a college education but the husband (wife) does not; the returns are only slightly higher, and the overtime trends are almost identical.

all marital statuses.[29] Second, the earnings return to college is higher for single females than for single males. The earnings return to college for single females has increased more than that for single males between 1946 and 1971 cohorts. Third, the earnings return to college is similar for married females and married males. Fourth, the earnings return to college is higher for divorced females than for divorced males.

We now analyze the case in which no change in earnings has occurred since 1946. The results are shown in Figure 16. Male and female attainment rates drop by 15.5 and 14.2 percentage points, respectively, by 1996. This indicates that the increasing returns to college in the labor market for those cohorts, as shown in Figure 15, have an important impact on college attainment for those cohorts.

The change of earnings has a larger effect on college attainment for females than that for males. This is due to the fact that over time the earnings return to college for single females has been increasing at a faster rate than that for single males. The gender reversal of college attainment occurs in the same year, however, as in the benchmark model. Thus the change in earnings over time cannot account for the reversal of the gender gap in college attainment.[30]

5.2.3 Marriage market

The next several simulations try to isolate the effects of changes in the marriage market on college attainments. First, we quantify the effects of declines in marriage and rises in divorce probabilities, keeping conditional marriage probabilities as in the data. Then, we show the effects of changes in conditional marriage probabilities, keeping single, marriage, and divorce probabilities as in the data.

[29] Using cross-sectional earnings or wages, many authors have documented recent increases in the earnings return to college (see, for example, Juhn, Murphy, and Pierce 1993; Katz and Murphy 1992; Card and DiNardo 2002; and Eckstein and Nagypál 2004). Our measure using lifetime earnings gives similar results.

[30] We also simulate a version that fixes the monetary costs of attending college at the 1946 cohort level, and the resulting attainments for both genders are very similar to the benchmark results.

Marriage probabilities Now we fix transition probabilities from single to married at the 1946 cohort level. Figure 17 shows that without decreases in marriage probabilities both males and females would reach higher college attainment in 1996. The gender reversal of college attainment occurs 4 years later than in the benchmark model.

The decrease in marriage probabilities decreases college attainment for both males and females. This can be explained by the differences in the total returns to education by marital status. As is shown in Figure 15, the earnings return to college in the labor market is higher for single females than for married females. However, married females receive an additional benefit from college by increasing their children's learning ability. Under our parameters, the return from children for married couples dominates their lower return in the labor market; thus, the returns to college increase with marriage probability. For males, the earnings return to college in the labor market is lower for single males than for married males. In addition, married males benefit from increasing their children's ability. Thus, the returns to college increase with marriage probability. As the marriage probability declines, returns to college decrease and so does college attainment.

The comparison also indicates that as marriage probabilities decline female college attainment decreases less than that of males. This occurs because single females receive a larger return to college in the labor market than do single males. Moreover, in our model, fathers do not enjoy their out-of-wedlock children's ability while mothers do. As a result, the decline in marriage probabilities decreases the returns to college for females less than those for males. Therefore, college attainment for females declines less than that for males.

Divorce probabilities Now we fix transition probabilities from married to divorced at the 1946 cohort level. Figure 18 shows that without change in the probabilities of divorce males would reach higher college attainment and females would reach lower college attainment in 1996. Females' college attainment would always be lower than that of males.

The increase in divorce probabilities decreases college attainment for males. This can be explained by differences in the earnings return to education by marital status. As is shown in Figure 15, the return to college is lower for divorced males than for married males. As the divorce probability increases, returns to college decrease and so does college attainment. The opposite happens for females: The return to college is higher for divorced females than for married females. As the divorce probability increases, returns to college increase and so does college attainment. The comparison also indicates that increase of divorce probabilities is the main force to account for the reversal of gender gap in college attainment. Without the increase of divorce probability, female college attainment never exceeds male attainment.

Marriage/single/divorce probabilities Now we fix both transition probabilities from single to married and transition probabilities from married to divorced at the 1946 cohort level. Figure 19 shows that without change in the probabilities of single, marriage, and divorce males would reach higher college attainment in 1996. Females' college attainment would change only slightly and would always be lower than that of males.

As is shown in the last two experiments above, both the increase of single and divorce probabilities decreases college attainment for males. Now in this experiment those two forces work together to reduce college attainment for males. An increase of single probabilities decreases college attainment for females, while an increase of divorce probabilities increases college attainment for females. Those two forces work in opposition to each other, and thus females' college attainment barely changes.

Conditional marriage probabilities Next we fix the conditional marriage probabilities at the level they were in 1946 and keep the marriage probabilities in the data. The results are shown in Figure 20. The college attainment in 1996 would be 3.3 percentage points lower for males. Therefore the change in conditional marriage probabilities plays a quantitatively minor role in accounting for the increase in college attainment for both genders.

The gender reversal of college attainment occurs 2 years earlier that in the benchmark model. The change of the marriage probability has a larger effect on college attainment for males than for females. This is in part due to the fact that over time the probability of marrying a college spouse for males has increased quite substantially, while the probability for females has barely changed. In our model, spousal education increases household income and children's human capital. In the benchmark, males over time benefit more from marrying college spouses than females do; thus, college attainment for males increases more than that for females.

Out-of-wedlock birth rate Figure 21 shows that without a rising out-of-wedlock birth rate females would reach a lower college attainment in 1996, while males keep the same college attainment. In our model, out-of-wedlock children do not affect males' education decisions. Although it is costly for single mothers to raise children, they value their children's learning ability. This gives females additional incentive to go to college.

6 Conclusion

We develop a dynamic model of college entry decision that incorporates intergenerational persistence on learning ability, marriage, and differential earnings by gender and marital status. Using this model, we study the quantitative effects of changes in relative earnings, changes in parental education, and changes in the marriage market on changes in college attainment by gender. We find that increases in parental education and relative earnings between college and high school persons increase college attainment for both genders. The rising divorce probabilities increase college attainment for females and decrease that for males, and thus are crucial in explaining the reversal of the gender gap in college attainment.

There are several directions in which this work can be extended. We assume marriage probabilities and earnings are exogenous. An extension that we wish to explore is the relationship among college attainment, marriage, and labor supply for both genders. Even though labor earnings are sacrificed, a parent who stays at home and takes care of children contributes to the household by increasing the learning ability of children. We plan to study these issues in future work.

7 Appendix

7.1 PSID sample

The PSID is a longitudinal survey of U.S. families and the individuals who make up those families. Approximately 4,800 U.S. families were sampled in 1968, and these families were reinterviewed annually until 1997. From 1997 onwards, PSID was changed to a biennial data collection and two major changes were made: a reduction of the core sample and the addition of a new sample of post-1968 immigrant families and their adult children.

We first find parents' education for the selected sample by linking parents and children from Individual Files (1968–2005). The PSID facilitates the intergenerational linkage by providing the parent's ID in the Individual Files. If a linkage cannot be found in Individual Files, we use 2003 Parent Identification Files to link an individual with his or her parents. If the above procedure fails to provide parents' education information, we find parents' education by using parents' and parents-in-law's education as reported by the head in Family Files. In 1974, questions were asked about how much education had been completed by the household head's parents and by the spouse's parents. In the later waves, these parental education questions were asked for new heads and spouses. By merging Individual Files with Family Files, we are able to find parents' education for those who were heads or spouses or siblings of the heads.

7.2 Estimation of wage

The model is estimated on the March CPS from 1964 to 2007. We restrict the sample to individuals who are between the ages of 18 and 65 who are not in the armed forces and not self-employed. To be consistent with the decision model, we restrict our attention to individuals who are either married or single (never married). Hourly wage is deflated to 2006 dollars using the CPI. Definitions of variables are given in Appendix section 7.2.2. We run separate probit wage selection and log wage regression for each gender in each year. The reduced-form probit selection results and estimated coefficients of the wage equations in 2007 are provided in Appendix sections 7.2.3 and 7.2.4.

7.2.1 Estimation procedure of wages

Consider the following wage function on a sample of working men and women:

$$\log w_i = X_i\beta + \mu_i,$$

where $\log w_i$ is the logarithm of hourly wage, and X is a vector of characteristics such as schooling and work experience. It is argued, however, that the sample of employed workers is not a random sample and that this selectivity might bias the coefficients. Formally, we can write down a participation equation

$$\begin{aligned} E_i &= 1 \text{ if } Z_i\gamma + \varepsilon_i \geq 0, \\ E_i &= 0 \text{ if } Z_i\gamma + \varepsilon_i < 0, \end{aligned}$$

where Z includes variables that predict whether or not a person works. Therefore, the probability of an individual working is

$$\Pr(E_i = 1) = \Pr(\varepsilon_i \geq -Z_i\gamma) = \Phi(\frac{Z_i\gamma}{\sigma}), \tag{12}$$

where σ_ε^2 is the variance of ε_i, and $\Phi(\cdot)$ is cumulative distribution function of the standard normal.

The selectivity problem is apparent by taking expectations of the wage function over the sample of employed workers:

$$E(\log w_i | E_i = 1, X_i) = X_i\beta + E\left(\mu_i | \varepsilon_i \geq -Z_i\gamma\right).$$

Supposing μ_i and ε_i are jointly normally distributed, let $\sigma_{\mu,\varepsilon}$ be the covariance between μ_i and ε_i. We can now write

$$E\left(\mu_i | \varepsilon_i \geq -Z_i\gamma\right) = \frac{\sigma_{\mu,\varepsilon}}{\sigma_\varepsilon}\frac{\phi(Z_i\gamma/\sigma)}{\Phi(Z_i\gamma/\sigma)},$$

where $\phi(\cdot)$ is the standard normal density. When $\sigma_{\mu,\varepsilon}$ is not zero, selectivity bias occurs. To estimate the potential wage consistently, we need to add the selection term (the inverse Mills ratio)

$$\frac{\phi(Z_i\gamma/\sigma)}{\Phi(Z_i\gamma/\sigma)} \equiv V_i \tag{13}$$

in the OLS regression as

$$\log w_i = X_i\beta + \alpha V_i + \eta_i.$$

7.2.2 Definitions of variables in X and Z

Age	Respondent's age
Age^2	Square of variable "Age"
HI	Dummy variable: 1 if respondent is a high school dropout
HG	Dummy variable: 1 if respondent is a high school graduate
SC	Dummy variable: 1 if respondent has some college education
CG	Dummy variable: 1 if respondent is a college graduate
Exp	Respondent's years of work experience
Exp^2	Square of variable Exp
Black	Dummy variable: 1 if respondent is black
Married	Dummy variable: 1 if respondent is married
Nchild	Number of own children in household
Nchlt5	Number of own children under age 5 in household
Northeast	Dummy variable: 1 if household is located in Northeast area
Midwest	Dummy variable: 1 if household is located in Midwest region
South	Dummy variable: 1 if household is located in South region
West	Dummy variable: 1 if household is located in West region
Metro	Dummy variable: 1 if household is located in a metropolitan area
Manager	Dummy variable: 1 if respondent is a manager or professional
Whitecollar	Dummy variable: 1 if respondent has white-collar occupation other than those in management
Bluecollar	Dummy variable: 1 if respondent has blue-collar occupation
V	See Equation (13)

7.2.3 Estimation results: probit selection

The reduced-form probit selection rule in equation (12) is estimated in each year for men and women. We estimate these probits year by year because some evidence shows that how individuals select themselves into the workforce has shifted over time (Mulligan and Rubinstein 2007). Table 2 presents estimated coefficients and asymptotic t-statistics of the reduced form participation probit for 2007.[31] Our findings are generally in accord with previous research. Specifically, we find that educational attainment has a positive and statistically significant impact on the probability of participation for both men and women. The probability of working increases in age at a decreasing rate for both men and women. Black men are less likely to participate than nonblacks. Men who are married or have children are more likely to participate than other men, even though the effect of the number of children is not statistically significant. Married women and women with children are less likely to participate.

7.2.4 Estimation results: wage equations

Estimated coefficients and asymptotic t-statistics of the wage equations in 2007 corrected for selections are found in Table 3. Estimated coefficients on education, experience, occupation dummies, race, and region dummies are similar to estimates from typical wage

[31] Estimates for other years are available from the authors.

	Males		Females	
Variable	Coefficient	t	Coefficient	t
Constant	-2.5929	-41.75	-2.6571	-43.10
HG	0.3134	15.67	0.5029	24.77
SC	0.3882	18.68	0.6520	32.05
CG	0.7044	31.09	0.8212	38.86
Age	0.1627	46.82	0.1448	41.89
Age^2	-0.0022	-52.07	-0.0018	-44.29
Black	-0.3328	-16.14	-0.0018	-0.10
Marry	0.4641	22.41	-0.0499	-2.91
Nchild	0.0396	4.82	-0.0800	-12.86
Nchlt5	0.0315	1.79	-0.2708	-22.21
No. of obs.	48,145		51,315	
-2 ln(likelihood ratio)	8285.05		5252.96	
χ^2 degree of freedom	9		9	

Table 2: Participation Selection Rules: Probit Analysis (CPS 2007)

equations found in the literature. College education attainments are generally more important for women's wage than for men's. Experience has more of a positive impact on men's wage than on women's.

	Males		Females	
Variable	Coefficient	t	Coefficient	t
Constant	1.5958	29.79	1.4525	31.64
HG	0.3278	22.01	0.2822	13.43
SC	0.4650	28.34	0.4704	20.27
CG	0.7743	36.96	0.8072	31.61
exp	0.0462	19.04	0.0336	19.29
exp^2	-0.0009	-13.81	-0.0006	-15.43
manager	0.3618	36.55	0.4974	38.16
white-collar	0.0099	1.09	0.1966	19.19
Midwest	-0.0713	-6.83	-0.0746	-6.61
South	-0.0829	-8.39	-0.0766	-7.09
West	-0.0445	-4.38	-0.0453	-4.01
metro	0.1150	12.74	0.1476	15.27
black	-0.1424	-9.38	-0.0298	-2.41
married	0.2748	16.84	0.0425	4.01
V	0.3131	5.08	0.2283	6.51
R^2	0.3026		0.2362	

Table 3: Estimates of wage equation: CPS 2007

Selectivity biases are particularly interesting. One would expect that individuals with higher wage potential should be more likely to participate in the labor force. The estimation results confirm that individuals who expect to earn more are more likely to participate in the labor force. The coefficients of V (defined in equation (13) in Appendix 7.2.1) are positive and statistically significant for both men and women. Therefore, observed wage patterns of men and women are higher than the population mean pattern would have been.

7.3 SMM Estimation Procedure

Let X_{ij} be the ith observation of the jth moment and denote N_j the number of individuals that comprise the jth moment. The sample moment is defined as

$$m_j = \frac{\sum_{i=1}^{N_j} X_{ij}}{N_j},$$

which is the average conditional attainment rate computed from PSID. The corresponding simulated moment is denoted by $m_j^S(\theta)$, which is computed based on the solution of the college entry decision model. Our task amounts to finding a parameter vector θ, which makes the model-simulated conditional attainment rates ($m_j^S(\theta)$) as close as possible to the empirical ones (m_j). The vector of moment conditions is

$$g(\theta)' = [m_1 - m_1^S(\theta), \cdots, m_j - m_j^S(\theta), \cdots, m_J - m_J^S(\theta)],$$

where J is the number of moments used and $J = 136$ (8 moments $\times$ 17 years). We minimize following objective function with respect to θ

$$L(\theta) = g(\theta)' W g(\theta), \tag{14}$$

where W is a weighting matrix.

Following Lee and Wolpin (2006), we make two assumptions in forming the weighting matrix W: (1) W is diagonal, (2) $E[g_j(\theta)^2] = \sigma_j^2/N_j$. We use a two-step procedure for computing the diagonal elements of W. First, we set $\sigma_j^2 = 1$ and weight each sample moment by N_j. We estimate θ by minimizing (14) and let $\widehat{\theta}$ be the first-stage estimate of θ. Second, we update σ_j^2 according to $\sigma_j^2 = g_j(\widehat{\theta})^2$. Then we weight each moment j by N_j/σ_j^2 and estimate θ again according to (14).

The variance-covariance matrix of the parameter estimates is given by $(A'WA)^{-1}$ where A is the matrix of the derivatives of the moments with respect to the parameters and W is the inverse of the variance-covariance matrix of the moments.

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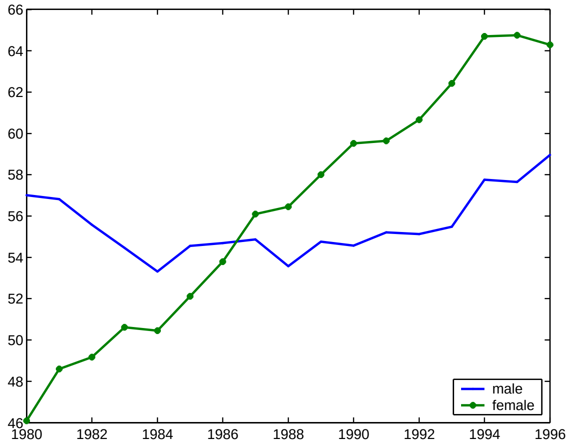


Figure 1: College attainment rates by age 34 among those aged 25–34.
Source: Authors' calculations from the PSID data files.

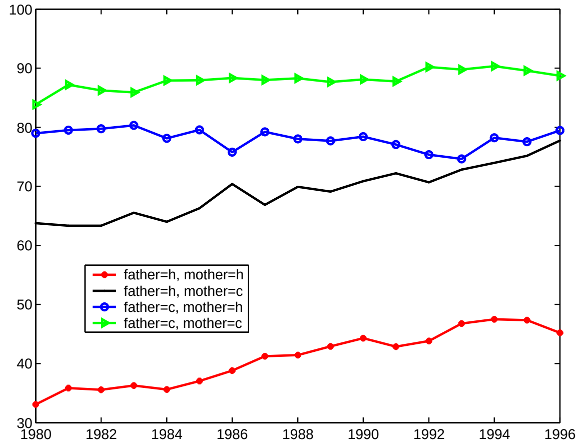


Figure 2: Female college attainment rates conditional on parental education.
Source: Authors' calculations from the PSID data files. h denotes high school and below, and c denotes some college and above.

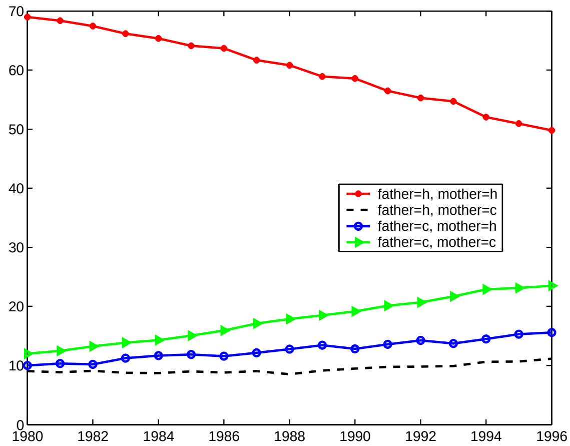
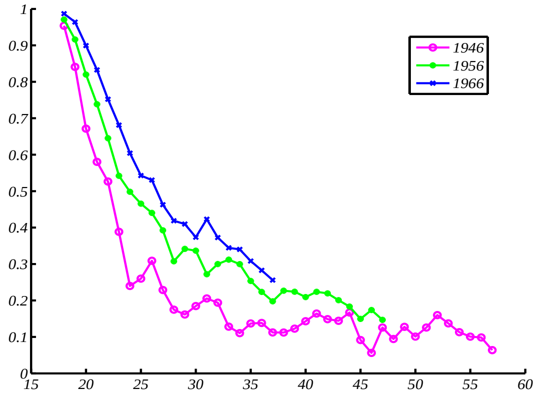
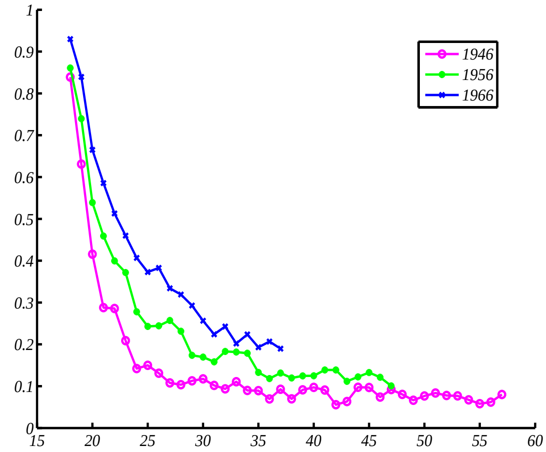


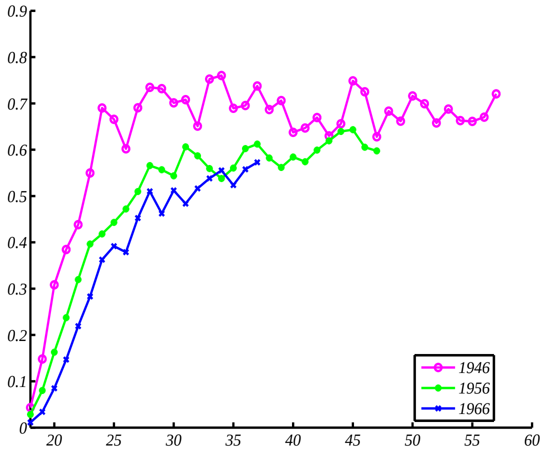
Figure 3: Parents' education distribution. Source: Authors' calculations from the PSID data files.



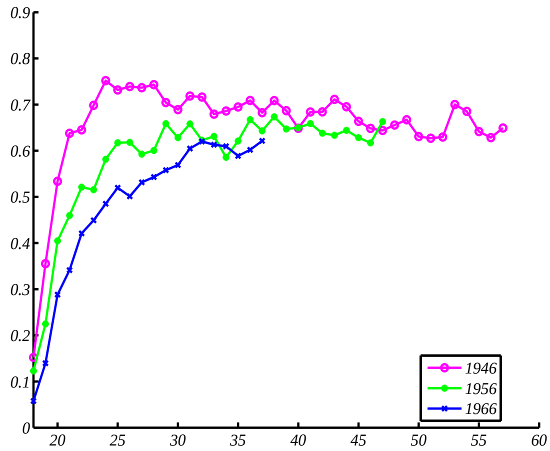
(a) Percent single for male



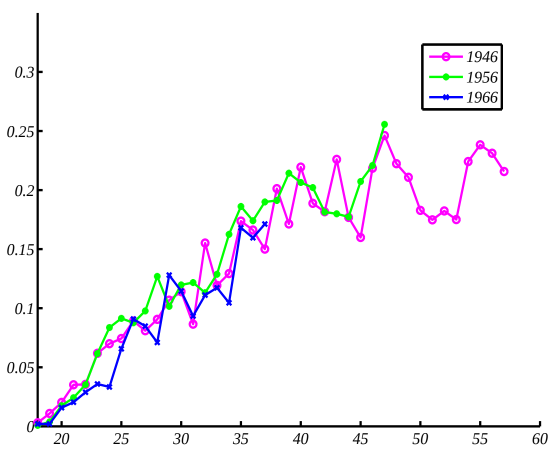
(b) Percent single for female



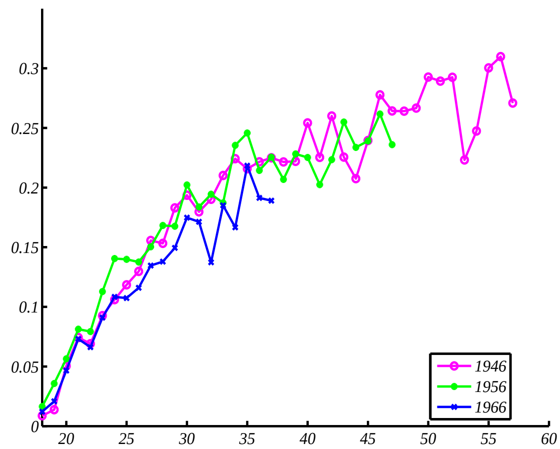
(c) Percent married for male



(d) Percent married for female

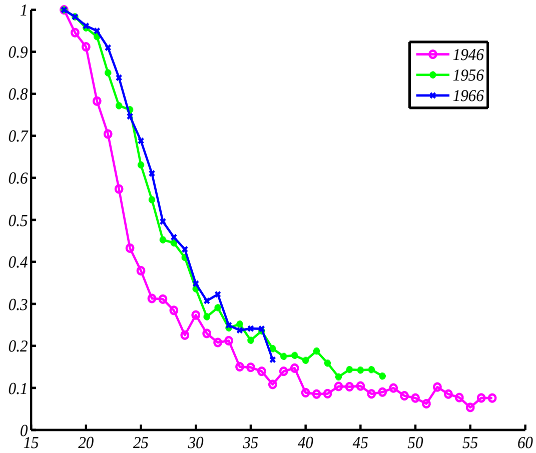


(e) Percent divorced for male

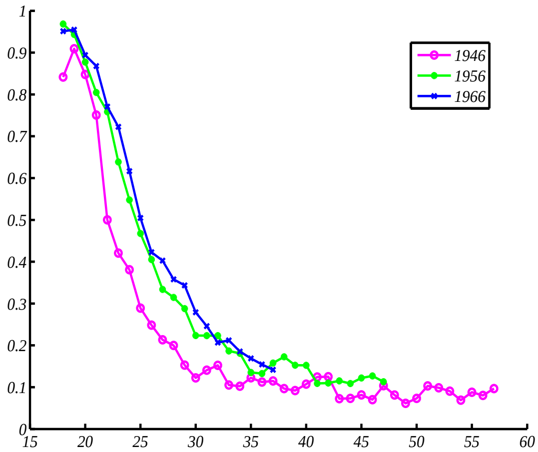


(f) Percent divorced for female

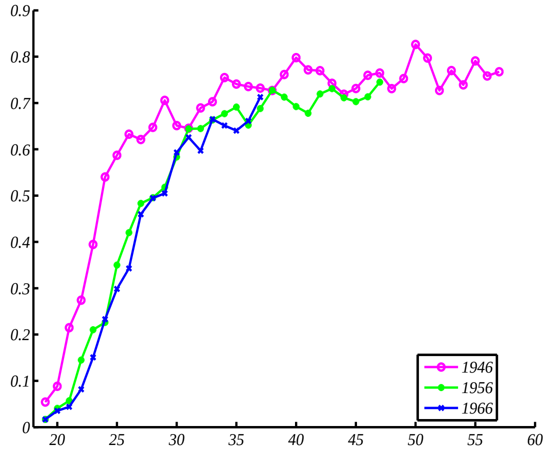
Figure 4: life-cycle profiles of marital status for high school



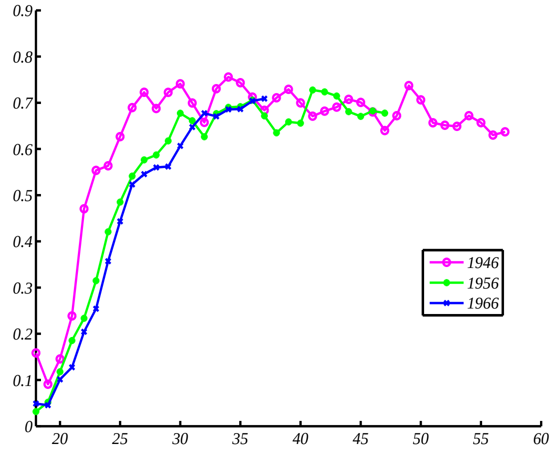
(a) Percent single for male



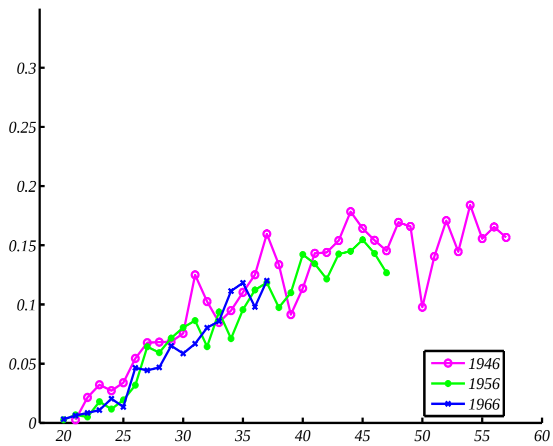
(b) Percent single for female



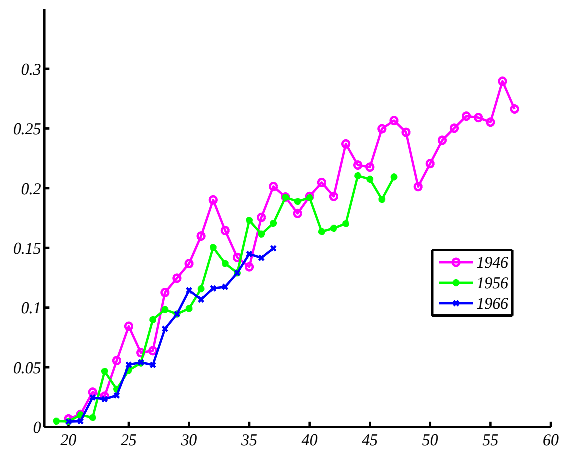
(c) Percent married for male



(d) Percent married for female



(e) Percent divorced for male



(f) Percent divorced for female

Figure 5: life-cycle profiles of marital status for college

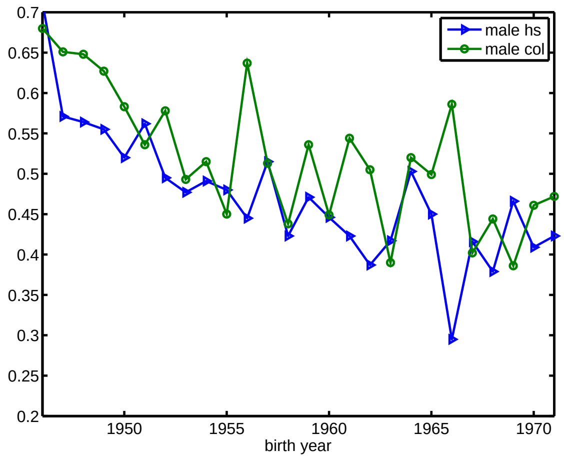


Figure 6: Marriage survival rate for males

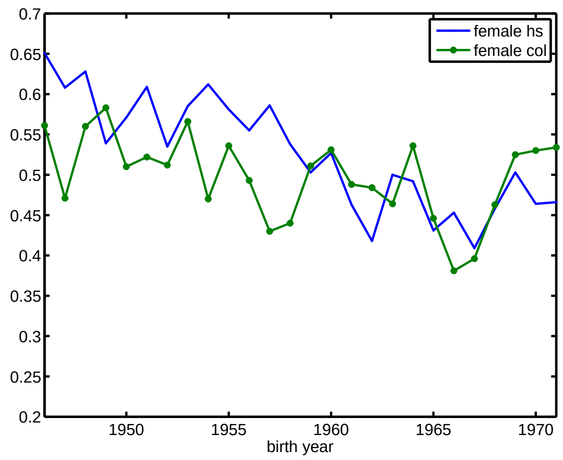


Figure 7: Marriage survival rate for females

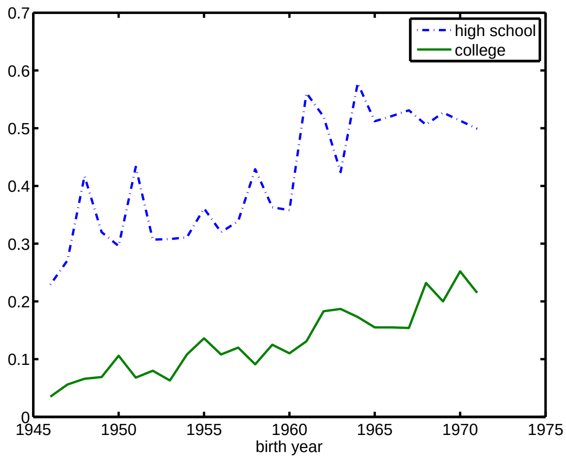
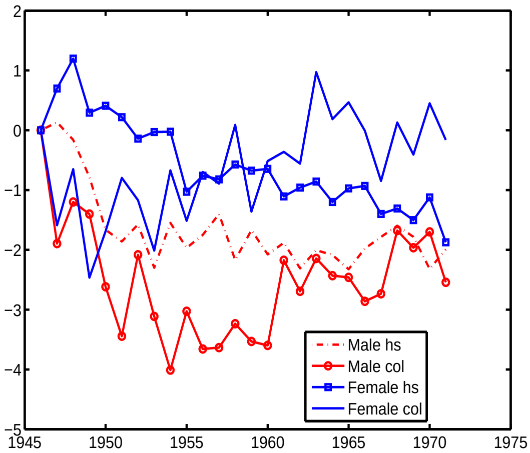
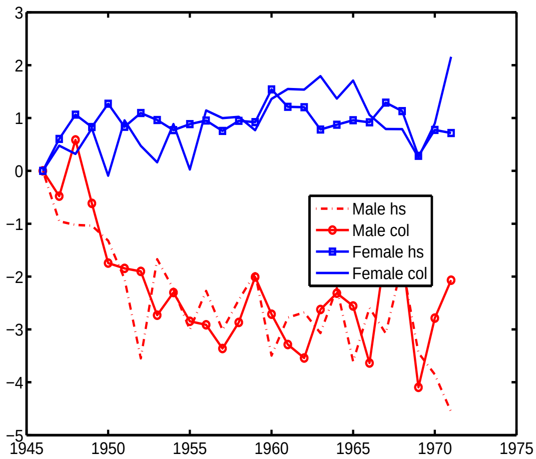


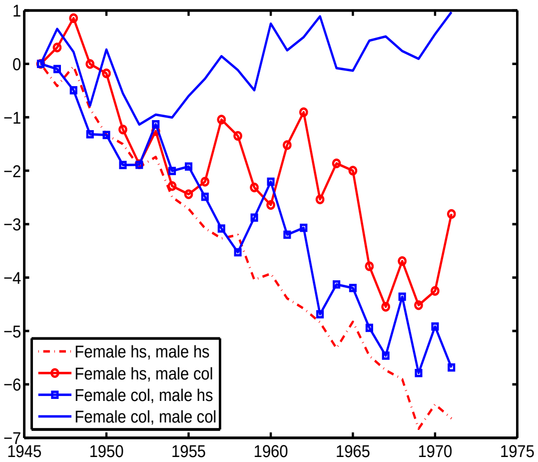
Figure 8: Probability of having out-of-wedlock children by age 30



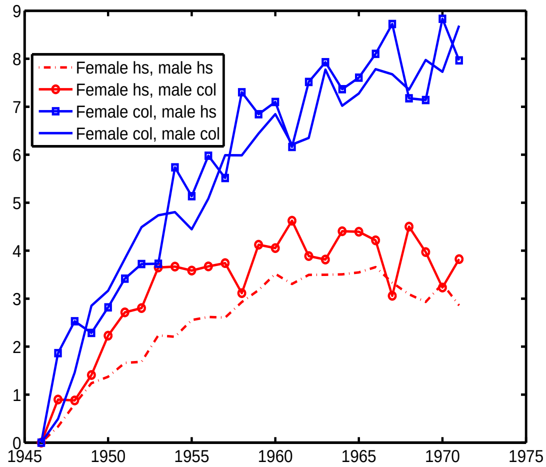
(a) Single



(b) Divorced

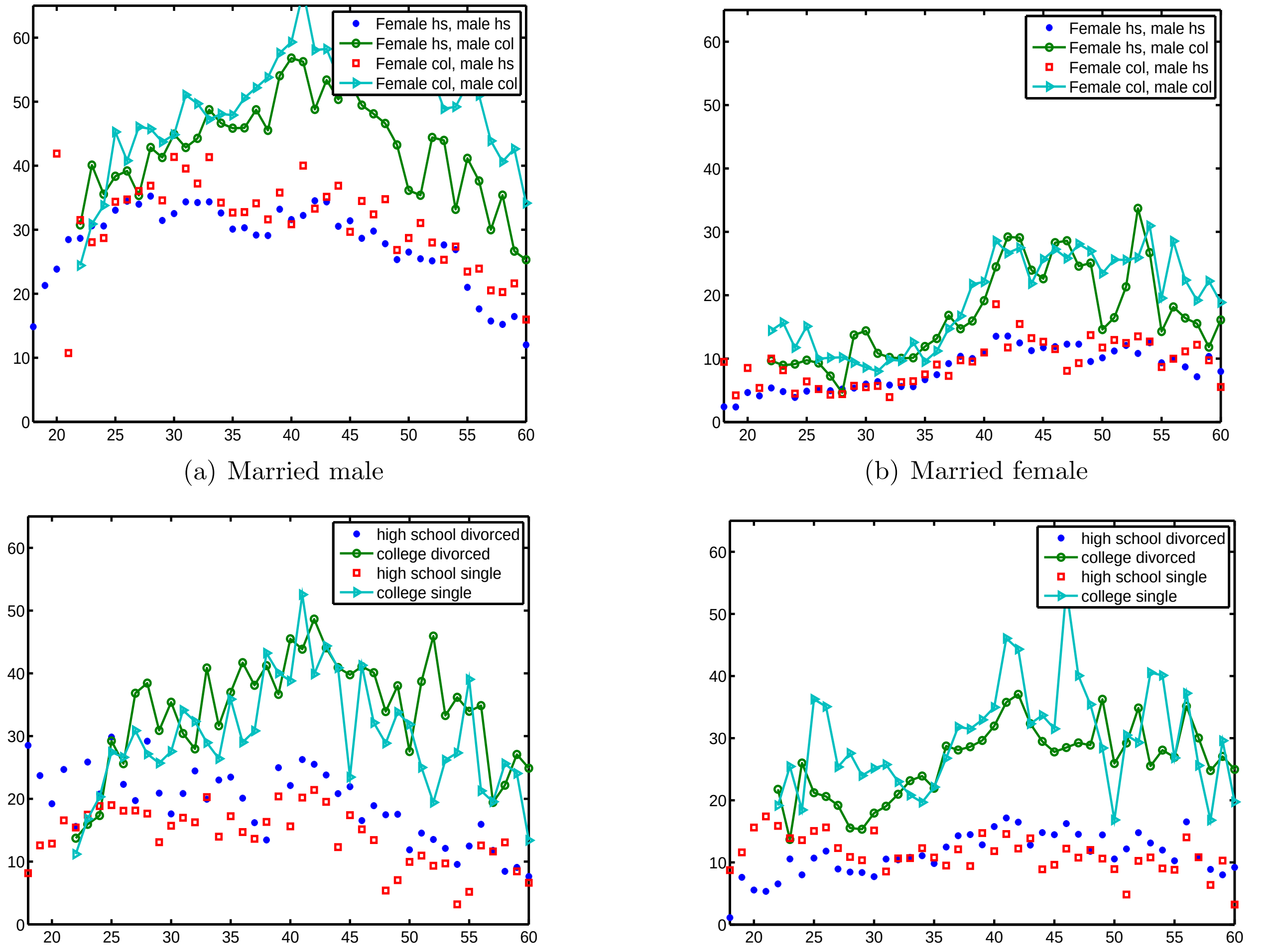


(c) Married male



(d) Married female

Figure 9: Cohort effects in earnings in 1000 dollars by marital status



(a) Married male (b) Married female

(c) Single and divorced male (d) Single and divorced female

Figure 10: life-cycle profiles of income for 1946 birth cohort in 1000 dollars, dotted: high school; solid: college

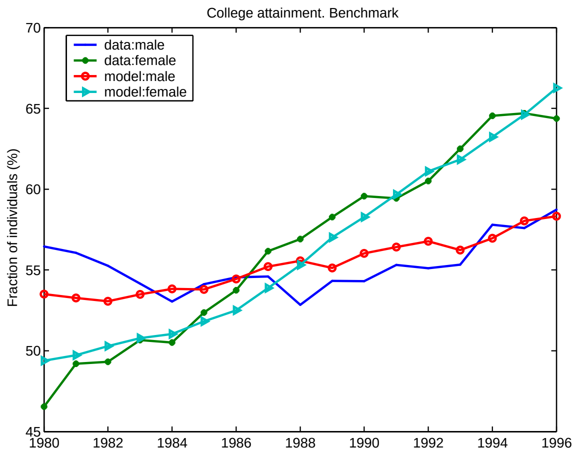


Figure 11: College attainment rates

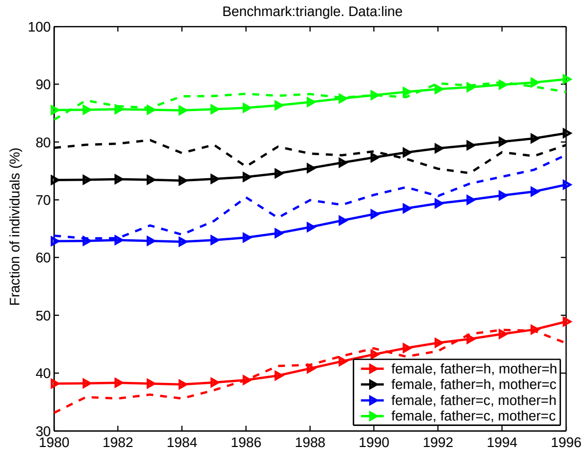


Figure 12: Female's college attainment rates (dashed line: data; solid line: model)

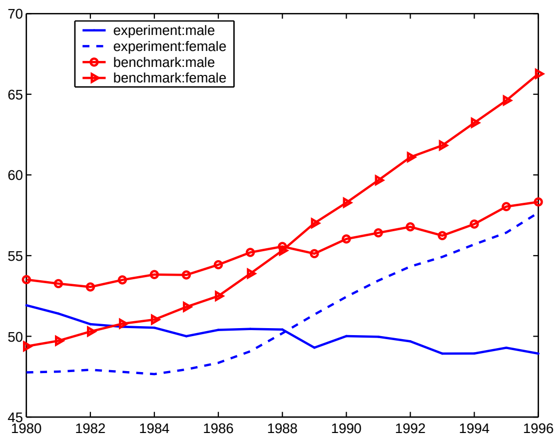


Figure 13: No change in parents' distribution since 1946

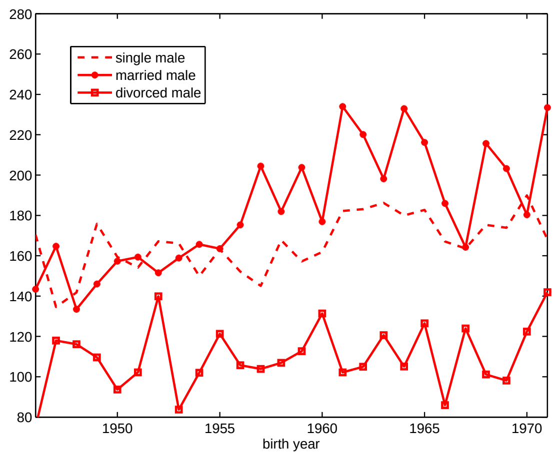


Figure 14: Earnings return to college for males by marital status in 1000 dollars

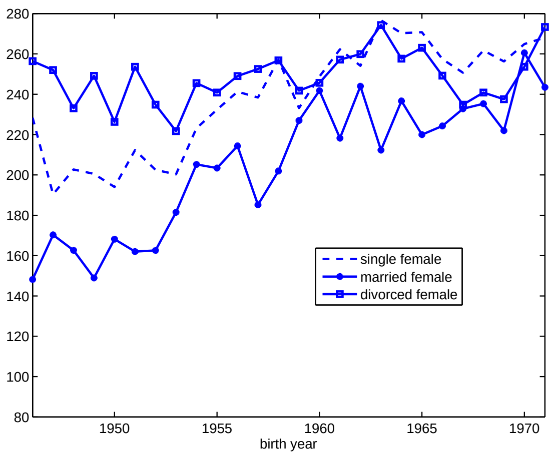


Figure 15: Earnings return to college for females by marital status in 1000 dollars

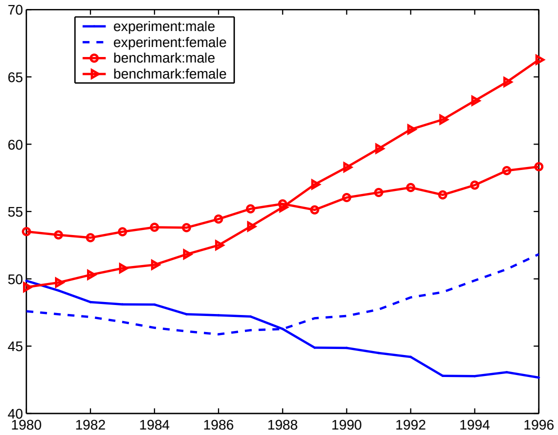


Figure 16: No change in earnings since 1946

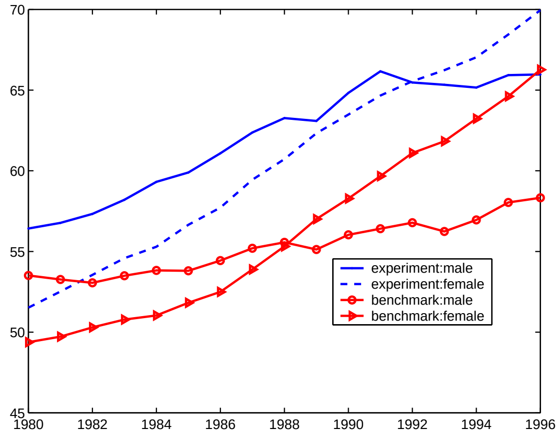


Figure 17: Marriage probabilities stay at 1946 cohort level

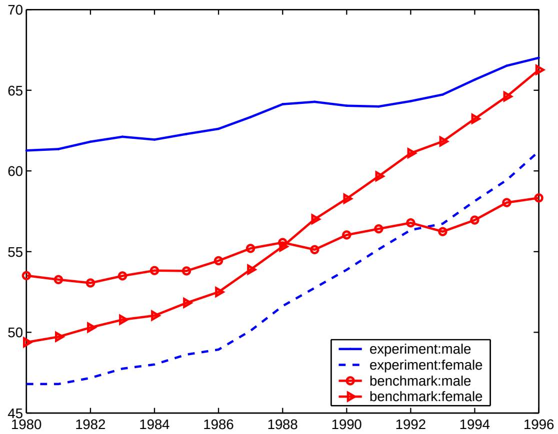


Figure 18: Divorce probabilities stay at 1946 cohort level

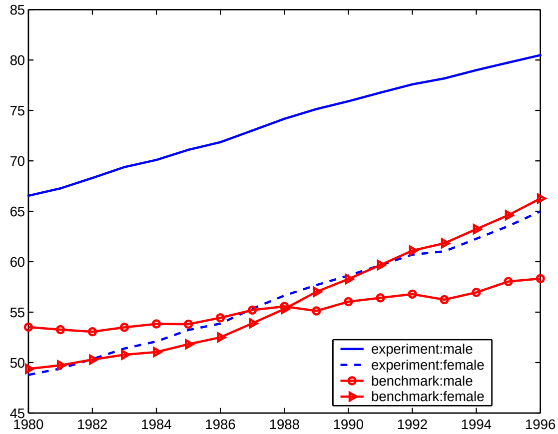


Figure 19: Marriage/single/divorce probabilities at 1946 cohort levels

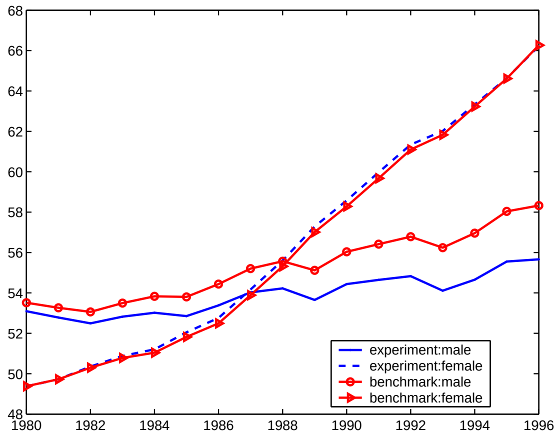


Figure 20: Conditional marriage probability stays at 1946 level

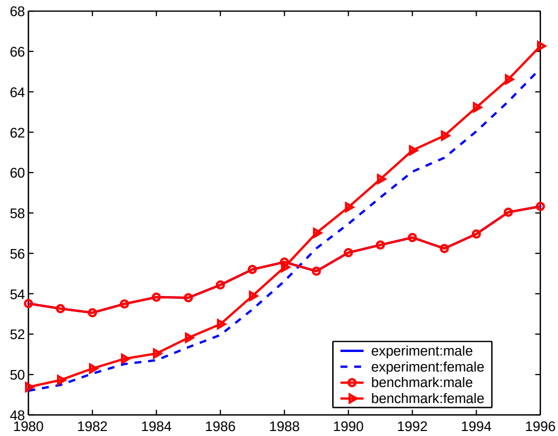


Figure 21: Out-of-wedlock birth rate stay at 1946 cohort level