ESSAYS IN LABOR ECONOMICS

By

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ABSTRACT

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This dissertation focuses on the interrelationship of socioeconomic outcomes of women. The first chapter of the dissertation explores the fragility of estimated effects of unilateral divorce laws on divorce rates and concludes that the impact of unilateral divorce laws remains unclear. I also make the methodological point that identification in differences-in-differences research becomes weaker in the presence of dynamics, especially in the presence of unit-specific time trends. The second chapter examines the plateau in U.S. women's labor force participation from the early 1990s to the present. The series of shift-share analyses shows that, for the most part, both the plateau and the earlier upward trend appeared within almost every category broken down by education, marital status, and child-rearing. In the third chapter, I look at the role of teenage women's anticipated future labor force attachment in explaining the upward trend in U.S. women's college-going. Combined with the trend towards higher work expectations of young women across birth cohorts, the results suggest that teenagers' future work expectations may account in part for the upward trends in women's college attendance and completion.

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Jongwoo Park, my love, provided abundant support throughout my time at Michigan State University. His love made all the difference.

Finally, I dedicate this dissertation to my parents.

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CHAPTER 1

THE FRAGILITY OF ESTIMATED EFFECTS OF UNILATERAL DIVORCE LAWS ON DIVORCE RATES[1]

1.1 Introduction

An influential *American Economic Review* article by Leora Friedberg (1998) estimated the impact on divorce rates from the adoption of unilateral divorce laws by most states during the 1970s. Using state/year data for 1968-1988, Friedberg estimated fixed-effects regressions of the annual per-capita divorce rate on a dummy variable for unilateral divorce with controls for state effects and year effects. In the simplest specification, she estimated almost exactly a zero effect of unilateral divorce, but in her preferred specifications that also controlled for linear or quadratic state-specific time trends, she estimated that adoption of unilateral divorce increased the divorce rate by more than 0.4 divorces per 1,000 population. This estimated impact seemed substantial relative to the sample mean divorce rate of 4.6, and Friedberg concluded (p. 608), "The move towards unilateral divorce accounted for 17 percent of the increase in divorce rates between 1968 and 1988."

In a subsequent *American Economic Review* article, Justin Wolfers (2006) questioned Friedberg's assumption that unilateral divorce shifted the divorce rate by a permanent constant. Instead, he posed the possibility that the impact might follow a distributed-lag pattern. Expanding the sample period to 1956-1988, he estimated regressions similar to Friedberg's except that he replaced her single dummy variable for unilateral divorce with a series of

[1] This chapter is co-authored with Gary Solon, Professor in Economics at Michigan State University.

dummies indicating that unilateral divorce had been in place for up to 2 years, 3-4 years, 5-6 years, …, 13-14 years, or at least 15 years. Based on the results, he concluded (p. 1802) "that the divorce rate rose sharply following the adoption of unilateral divorce laws, but that this rise was reversed within about a decade." In the present note, we check the robustness of Wolfers's estimates to some reasonable variations in estimation method and functional form.[2]

1.2 Robustness Checks

1.2.A Weighted and Ordinary Least Squares Estimates of Linear-in-Levels Specifications

Like Friedberg, Wolfers estimated his regressions by weighted least squares (WLS), weighting each state/year observation by the state's population. The results are shown in the first columns of our tables 1.1, 1.2, and 1.3, respectively for models with no state-specific trends, linear trends, and quadratic trends. Thanks to Wolfers's generous posting of his data and code on his website, it was easy to obtain exact replications of his results. In all three specifications, unilateral divorce is estimated to increase the divorce rate initially by about 0.3 divorces per 1,000 population. The basis for Wolfers's conclusion about dynamics is that the impact appears to remain substantial throughout the first eight years since adoption, but beyond eight years there is little evidence of a substantial positive impact.

Like Friedberg, Wolfers estimated his standard errors (shown in parentheses in the first columns of our tables 1.1-1.3) on the assumption that the error term in each weighted regression is homoskedastic and serially uncorrelated. Setting aside the homoskedasticity assumption for

[2] It is understandable that Wolfers hewed closely to Friedberg's estimation method and functional form in an effort to explore the role of dynamics while holding these other factors constant.

the moment, our analysis of the residuals from Wolfers's regressions shows overwhelming evidence of strong serial correlation. First-order autoregressions of the WLS residuals show estimated autocorrelations of 0.93, 0.94, and 0.89 respectively for the three specifications. Furthermore, the estimated autocorrelations remain substantial at higher orders. For example, the respective estimates of the fifth-order autocorrelations are 0.73, 0.73, and 0.53.[3] Obviously, neglect of such strong serial correlation could lead to serious bias in the estimation of standard errors. We therefore have used Stata's *cluster* option to implement Arellano's (1987) method of correcting standard error estimates for both serial correlation and heteroskedasticity. The resulting standard error estimates, shown in brackets in the first columns of our tables 1-3, are dramatically larger than those reported by Wolfers. Once it is recognized that the coefficient estimates may be much less precise than they previously appeared, the statistical significance of the coefficient estimates for lags up to eight years becomes less compelling.[4]

An even more striking pattern emerges when we reconsider whether weighting by population is an appropriate correction for heteroskedasticity. The common practice of weighting by population leads to efficient coefficient estimation under the very strong assumption that the error terms for individuals within the state are homoskedastic and independent of each other. As explained in detail by Dickens (1990), however, in the likely case that the individual error terms are positively correlated because they share a common state-level

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[3] As shown in Nickell (1981) and Solon (1984), these estimates are biased downward by the incidental-parameters problem in fixed-effects estimation. With a time-series length of 33 years, however, the bias is small.

[4] This same point is made by Vogelsang (2008) with respect to the model with no state-specific trends. We say "may be less precise" rather than "are less precise" because Vogelsang goes on to find that, when he also applies his method for correcting for spatial correlation (correlation among error terms across states), the standard error estimates come back down. In the end, Wolfers's assumption of homoskedasticity and no serial correlation clearly is incorrect, but it is not so clear what the precision of the coefficient estimates really is.

error component, the original (unweighted) state-average error terms may be nearly homoskedastic, and weighting by population size may *induce* a substantial heteroskedasticity problem. In this case, ordinary least squares (OLS) may be considerably more efficient than WLS. At minimum, then, it is good practice to report coefficient estimates from both methods, along with heteroskedasticity-robust standard error estimates. Also, as emphasized by DuMouchel and Duncan (1983), since both estimators are consistent if the model otherwise is correctly specified, discrepancies between WLS and OLS estimates can be a basis for testing model specification.

In the second columns of tables 1.1-1.3, we report the results from estimating by OLS instead of WLS. Where the WLS coefficient estimates for lags up to eight years had appeared substantially positive, this is much less true for the OLS coefficient estimates in all three specifications. This same finding has been reported by Dröes and van Lamoen (2010). Furthermore, in accordance with Dickens's critique of population-weighting, although the robust standard error estimates (shown in brackets) appear larger for OLS than for WLS in table 1, this is less clear for the specifications in tables 1.2 and 1.3.

In any case, the substantial discrepancies between the WLS and OLS estimates are suggestive of model misspecification. We will take up the issue of functional form in the next sub-section. Here, we will close by noting another potential specification issue – the assumption that the vector of distributed-lag coefficients is homogeneous across states may be incorrect. If the impact of unilateral divorce is heterogeneous – i.e., if it interacts with other state characteristics – then WLS and OLS estimates that do not explicitly account for those interactions may identify different averages of the heterogeneous effects. By design, WLS places more weight on the more populous states. If unilateral divorce tends to have larger effects in those states, then the WLS estimates will tend to be larger than the OLS estimates. This possibility is suggested by the results in the third columns of tables 1.1-1.3, which redo WLS with California omitted from the sample. The coefficient estimates are much smaller than the full-sample WLS results in the first columns. The discrepancy between the estimates in columns 1 and 2 may arise partly because California looms large in the full-sample WLS estimates, but is weighted the same as any other state in the OLS estimates.[5]

1.2.B Alternative Functional Forms

All the regression models estimated by Wolfers, as well as Friedberg, are linear models for the level of the divorce rate. This is not the only possible functional specification, and arguably not the most natural. For example, given that the divorce rate is always positive, linear models for the logarithm of the divorce rate are an obvious alternative. Indeed, since the divorce rate is a fraction, linear models for the logit of the divorce rate (i.e., the log odds of divorcing) also seem appealing. In practice, though, there is almost no difference in the present case between the log and the logit. With the divorce rate R expressed per capita, the logit of R is $\log[R/(1-R)] = \log R - \log(1-R)$, which is almost identical to $\log R$ because $\log(1-R)$ is so close to zero. In the rest of tables 1.1-1.3, we report results for only the log of R because the logit results are virtually identical.

The last two columns in tables 1.1-1.3 report both WLS and OLS estimates of models in which the dependent variable is the divorce rate in log rather than level form. We think this sensitivity analysis is our most striking of all. Once the divorce rate is in log form, neither the

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[5] As noted by Friedberg (1998, footnote 20), Nevada is a major outlier with respect to divorce rates. We have found, however, that, with state fixed effects accounted for, excluding/including Nevada is much less consequential for the gap between WLS and OLS estimates than is the treatment of California.

WLS nor OLS estimates show any indication of a positive effect of unilateral divorce on divorce rates in any of the three specifications. Whereas Friedberg and Wolfers differed over whether the positive impact of unilateral divorce on divorce rates is permanent or temporary, the log results show no positive impact at all. The fragility of the previously published estimates does not prove those estimates are wrong. Rather, the extreme sensitivity of the results to functional form assumptions, as well as estimation methods, leaves unclear what the true impact of unilateral divorce has been.[6]

Three other points about functional form are worth noting. First, the discrepancy between the levels results and the log (and logit) results is a vivid counter-example to the frequent claim that functional form assumptions are of little consequence in the estimation of models for binary, fractional, or otherwise limited dependent variables.[7] Second, compared to the levels specifications, the log specifications show less discrepancy between the WLS and OLS results. Third, with regard to Dickens's analysis of weighting, the robust standard error estimates for the log specifications suggest that, for those specifications, OLS is considerably more efficient than WLS.

1.3. Summary and Discussion

[6] It may be worth adding that none of the estimates in our article, Wolfers's, or Friedberg's indicates that the adoption of unilateral divorce accounts for most of the rise in divorce rates during the 1970s. As Stevenson and Wolfers (2007, p. 46) put the point, "Despite apparent conflict in this literature, it is worth emphasizing the point of substantial agreement: each of these authors finds that liberalized divorce laws had at most a small effect on divorce rates, and these reforms explain very little of the rise in divorce over the past half century."

[7] For example, a lengthy passage on this topic in Angrist and Pischke (2009) concludes (p. 107), "The upshot of this discussion is that while a nonlinear model may fit the CEF [conditional expectation function] for LDVs [limited dependent variables] more closely than a linear model, when it comes to marginal effects, this probably matters little. This optimistic conclusion is not a theorem, but, as in the empirical example here, it seems to be fairly robustly true."

Our main substantive finding is that estimates of the impact of unilateral divorce laws on divorce rates are highly sensitive to variations in model specification and estimation method. We do not take this to mean that the conclusions in Wolfers (2006) are necessarily wrong. We take it to mean that the true impact of unilateral divorce laws remains unclear.

In addition, the Wolfers article appropriately emphasized broader methodological lessons. Having found that Friedberg's (1998) results seemed sensitive to allowance for dynamic response, Wolfers concluded (p. 1802),

> A major difficulty in difference-in-difference analyses involves separating out preexisting trends from the dynamic effects of a policy shock. Her approach appears to confound the two. This problem – that state-specific trends may pick up the effects of a policy and not just preexisting trends – is quite general. Slight modifications to standard procedures yield more directly interpretable results.

We fully agree that, when the impact of a policy change is not necessarily immediate and constant, researchers should follow Wolfers's example of exploring the dynamics of the response. But the results we have reported suggest an additional cautionary note. The differences-in-differences research design with unit-specific time trends is essentially a type of regression discontinuity design, with time as the "running variable." As in other regression discontinuity designs, identification is based on the appearance and size of a "jump" in the dependent variable at the discontinuity point – in this case, the time at which the new policy is adopted. When the shift in the dependent variable may vary with the length of time since the policy change, and especially when that complication is accompanied by other differences across states in time trends, the sharpness of the identification strategy suffers. In such cases, we are asking the data a much more nuanced question than just whether the dependent variable series showed a constant discrete shift at the moment of policy adoption, and sometimes the answer will be unclear. In that broader methodological light, it may not be so surprising that the evidence on the impact of unilateral divorce laws is ambiguous.

Table 1.1 Estimated Effects of Unilateral Divorce Laws with No Controls for State-Specific Time Trends

	(1)	(2)	(3)	(4)	(5)
Dependent variable:	Divorce rate	Divorce rate	Divorce rate	Log of divorce rate	Log of divorce rate
Estimation method:	WLS	OLS	WLS excluding California	WLS	OLS
First 2 years	0.267 (0.085) [0.188]	-0.219 (0.192) [0.292]	0.071 (0.093) [0.131]	-0.136 (0.033) [0.114]	-0.091 (0.028) [0.051]
Years 3-4	0.210 (0.085) [0.159]	-0.273 (0.194) [0.423]	0.123 (0.093) [0.168]	-0.206 (0.033) [0.132]	-0.114 (0.029) [0.065]
Years 5-6	0.164 (0.085) [0.171]	-0.425 (0.198) [0.490]	0.097 (0.092) [0.184]	-0.249 (0.033) [0.144]	-0.152 (0.029) [0.076]
Years 7-8	0.158 (0.084) [0.174]	-0.452 (0.200) [0.477]	0.103 (0.091) [0.190]	-0.270 (0.032) [0.151]	-0.171 (0.029) [0.081]
Years 9-10	-0.121 (0.084) [0.163]	-0.703 (0.203) [0.479]	-0.144 (0.090) [0.178]	-0.331 (0.032) [0.157]	-0.220 (0.030) [0.083]
Years 11-12	-0.324 (0.083) [0.180]	-0.741 (0.203) [0.503]	-0.262 (0.090) [0.187]	-0.376 (0.032) [0.165]	-0.234 (0.030) [0.085]
Years 13-14	-0.461 (0.084) [0.199]	-0.845 (0.212) [0.544]	-0.390 (0.091) [0.211]	-0.419 (0.032) [0.170]	-0.265 (0.031) [0.088]
Years 15+	-0.507 (0.080) [0.233]	-0.776 (0.208) [0.472]	-0.348 (0.090) [0.224]	-0.493 (0.031) [0.188]	-0.326 (0.031) [0.098]

Notes: Divorce rate is number of divorces per 1,000 persons by state and year. Standard error estimates in parentheses assume homoskedasticity and serial non-correlation. Standard error estimates in brackets are robust to heteroskedasticity and serial correlation. All regressions include controls for year and state fixed effects.

Table 1.2 Estimated Effects of Unilateral Divorce Laws with Controls for Linear State-Specific Time Trends

	(1)	(2)	(3)	(4)	(5)
Dependent variable:	Divorce rate	Divorce rate	Divorce rate	Log of divorce rate	Log of divorce rate
Estimation method:	WLS	OLS	WLS excluding California	WLS	OLS
First 2 years	0.342 (0.062) [0.196]	0.141 (0.096) [0.112]	0.104 (0.066) [0.095]	-0.022 (0.020) [0.063]	-0.017 (0.018) [0.026]
Years 3-4	0.319 (0.070) [0.154]	0.211 (0.107) [0.120]	0.170 (0.073) [0.133]	-0.049 (0.023) [0.063]	-0.014 (0.020) [0.031]
Years 5-6	0.300 (0.077) [0.174]	0.177 (0.121) [0.180]	0.152 (0.080) [0.153]	-0.051 (0.025) [0.064]	-0.022 (0.023) [0.034]
Years 7-8	0.322 (0.084) [0.195]	0.250 (0.132) [0.283]	0.172 (0.087) [0.177]	-0.033 (0.028) [0.065]	-0.013 (0.025) [0.039]
Years 9-10	0.081 (0.091) [0.206]	0.133 (0.143) [0.370]	-0.049 (0.094) [0.202]	-0.052 (0.030) [0.067]	-0.030 (0.027) [0.046]
Years 11-12	-0.102 (0.099) [0.223]	0.144 (0.154) [0.424]	-0.164 (0.101) [0.240]	-0.051 (0.032) [0.074]	-0.015 (0.029) [0.052]
Years 13-14	-0.202 (0.107) [0.236]	0.210 (0.168) [0.484]	-0.270 (0.109) [0.248]	-0.043 (0.035) [0.077]	-0.005 (0.032) [0.060]
Years 15+	-0.210 (0.119) [0.263]	0.311 (0.187) [0.540]	-0.246 (0.121) [0.275]	0.006 (0.039) [0.084]	0.026 (0.035) [0.073]

Notes: Divorce rate is number of divorces per 1,000 persons by state and year. Standard error estimates in parentheses assume homoskedasticity and serial non-correlation. Standard error estimates in brackets are robust to heteroskedasticity and serial correlation. All regressions include controls for year and state fixed effects.

Table 1.3 Estimated Effects of Unilateral Divorce Laws with Controls for Quadratic State-Specific Time Trends

	(1)	(2)	(3)	(4)	(5)
Dependent variable:	Divorce rate	Divorce rate	Divorce rate	Log of divorce rate	Log of divorce rate
Estimation method:	WLS	OLS	WLS excluding California	WLS	OLS
First 2 years	0.302 (0.054) [0.169]	0.050 (0.075) [0.089]	0.110 (0.059) [0.093]	-0.021 (0.020) [0.053]	-0.015 (0.018) [0.021]
Years 3-4	0.289 (0.065) [0.127]	0.062 (0.092) [0.143]	0.198 (0.072) [0.131]	-0.047 (0.025) [0.054]	-0.014 (0.022) [0.026]
Years 5-6	0.291 (0.079) [0.141]	-0.036 (0.116) [0.179]	0.203 (0.090) [0.160]	-0.046 (0.030) [0.056]	-0.022 (0.028) [0.033]
Years 7-8	0.351 (0.097) [0.165]	-0.026 (0.144) [0.205]	0.257 (0.111) [0.202]	-0.019 (0.037) [0.061]	-0.008 (0.035) [0.043]
Years 9-10	0.161 (0.117) [0.172]	-0.210 (0.177) [0.240]	0.076 (0.137) [0.232]	-0.028 (0.045) [0.069]	-0.018 (0.043) [0.056]
Years 11-12	0.047 (0.142) [0.190]	-0.270 (0.215) [0.278]	0.006 (0.167) [0.280]	-0.014 (0.054) [0.088]	0.005 (0.052) [0.070]
Years 13-14	0.031 (0.167) [0.211]	-0.289 (0.257) [0.328]	-0.047 (0.199) [0.310]	0.011 (0.063) [0.098]	0.029 (0.063) [0.086]
Years 15+	0.251 (0.205) [0.261]	-0.226 (0.317) [0.407]	0.074 (0.242) [0.341]	0.104 (0.078) [0.113]	0.087 (0.077) [0.107]

Notes: Divorce rate is number of divorces per 1,000 persons by state and year. Standard error estimates in parentheses assume homoskedasticity and serial non-correlation. Standard error estimates in brackets are robust to heteroskedasticity and serial correlation. All regressions include controls for year and state fixed effects.

CHAPTER 2

THE PLATEAU IN U.S. WOMEN'S LABOR FORCE PARTICIPATION: A COHORT ANALYSIS

2.1 Introduction

One of the most remarkable changes in the U.S. labor market in the twentieth century was a massive increase in the rate of female labor force participation (FLFP). The FLFP rate was only 17 percent in 1890 (Goldin 2000, Table 10.1), but had more than tripled by a century later, reaching 59 percent in 1994 (Figure 2.1). This trend seems even more dramatic when contrasted with the gradual decline in men's labor force participation.[1] Since 1994, however, an important new pattern has emerged: after many decades of trending upward, the FLFP rate has leveled off at around 60 percent.[2] This recent stagnation in the labor force involvement of U.S. women is rare even in world economies (U.S. Bureau of Labor Statistics, table of women's labor force participation rates, selective countries, 1970-2009, available at http://www.bls.gov/spotlight/2011/women/).

The end of the upward trend in U.S. women's labor force participation is an important and puzzling phenomenon. For example, labor economists traditionally have described the long-running rise in women's labor force participation as a movement along a positively sloped labor

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[1] See Juhn and Potter (2006) and Fallick and Pingle (2007) for the trend in U.S. men's labor force participation.

[2] Although the 1994 Current Population Survey redesign has caused changes in the measurement of many of the statistics derived from the CPS (Polivka and Miller, 1998), trends in the FLFP are not sensitive to the redesign. All estimates presented in this paper are based on data unadjusted for the1994 CPS redesign. Changes in the percent institutionalized also are not an issue because the trend in the fraction of women institutionalized was stable over time. My calculation using U.S. Censuses shows that the percentages of women institutionalized (correctional and mental institutions and institutions for the elderly, handicapped, and poor) in 1980, 1990 and 2000 are 1.37, 1.49, and 1.34, respectively.

supply curve in response to real wage growth. The plateau in women's labor force participation, however, has occurred despite continued real wage growth (Eckstein and Lifshitz, 2011). A simple labor supply story therefore cannot account for the trend shift unless it involves a credible explanation for why the labor supply curve for women switched from quite positively sloped to inelastic.

This paper uses March Current Population Survey data from 1968 to 2010 to assemble a clear and detailed catalog of facts about the plateau in U.S. women's labor force participation. Following and extending earlier work by Goldin (2006) and Fallick and Pingle (2007),[3] I find that a great deal of what has happened to FLFP can be succinctly summarized in terms of trends across birth cohorts. In addition to providing a more up-to-date cohort analysis, I conduct shift-share analyses that document within-group trends and composition effects based on disaggregations by educational attainment, marital status, and child care responsibilities.[4] In the end, I deliver a constellation of evidence about the trends in women's labor force participation and related socioeconomic outcomes that I hope will stimulate and guide further research into the causes of the shifting trends.

Although numerous recent studies have analyzed FLFP in the United States, these studies are less comprehensive than mine in that they cover shorter time periods, are restricted to particular subpopulations, or provide no shift-share analyses using related socioeconomic outcomes. For example, Boushey (2005) and Hoffman (2009) find that the negative effects of

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[3] Aaronson et al. (2006) and Percheski (2008) are other precedents.

[4] Race is also an important demographic variable that may interact with women's labor supply, but it plays a minor role in explaining the plateau. Since the early 1990s, the FLFP rate among white women has shown a leveling-off, and the rate among black women has displayed a very slight decline. My sample includes all race and ethnic groups. Results in this paper still hold when the population is limited to white women only.

children on labor force participation for women aged 25 to 44 fell over time in the Current Population Survey's Outgoing Rotation Group data for 1984, 1989, 1993, 2000, and 2004. Percheski (2008) investigates cohort differences in female employment rates using data from decennial Censuses for 1960 through 2000 and the American Community Survey for 2005, but she restricts her samples to professional women. Goldin (2006) explores how women's employment, education, and family status have transformed together using various data sources, with an emphasis on the period before the 2000s. Her conjecture that the plateau in FLFP was a temporary phenomenon due to the recession of the early 2000s is testable (and refuted) by my updating of the evidence through 2010. Macunovich (2010) presents the change in FLFP by socioeconomic subgroup over time using March CPS data for 1976 through 2009, but does not look at compositional changes of the subgroups. Fallick and Pingle (2007) do decompose the changes in male and female labor force participation rates over time into group-share and within-group changes, but the subgroups for their decompositions are based only on age.

My series of shift-share analyses uniquely provides the quantitative connection between FLFP and related socioeconomic outcomes of women. It shows that both the rising FLFP and the plateau can be explained mostly by trends within almost every group – in particular, within every education and child-rearing group, and within the ever-married group.

2.2 Data

My data come from the Integrated Public Use Microdata Series (IPUMS) version of the March Current Population Surveys (CPS) for 1968 through 2010. I restrict my sample to noninstitutionalized civilian women aged 16 – 64.[5] I do not include observations before 1968 because a variable "MOMLOC," which is used to link mothers to their own children, is not available for those observations. Throughout this paper, I use single-year birth cohorts and five-year age bins (or a four-year age bin for the youngest age group). Thus, each cohort includes women born in the reference year, and the whole sample is divided into ten age groups, from age group 16-19 to 60-64. Birth year is calculated as calendar year minus age. In order to retain a sufficient sample size for each age-cohort group, the figures are based on the cohorts 1932-1982. To compute averages for each subpopulation, I use CPS-provided sampling weights. Under the simplifying approximation that the covariance of sampling errors between any two age-cohort groups is zero, most of my standard error estimates are less than one percent of the corresponding estimates. Therefore, I do not clutter my figures with confidence bands.

2.3 Cohort shifts in labor force participation and related socioeconomic outcomes of women

2.3.A Birth cohort and FLFP

Figure 2.2.A displays the life-cycle trajectories of labor force participation of women. The figure presents only nine of the cohorts for expositional simplicity. (The trajectory for the 1982 cohort is hardly distinguishable from those for the cohorts 1964-1976 due to their overlapping.) The choice of the cohorts is arbitrary and has no bearing on the conclusions. For the cohorts up through the 1952 birth year, the life-cycle trajectories are strongly inverse-U-shaped, with labor force participation peaking in the late forties. For more recent cohorts, the

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[5] The exception is Figure 2.1, which is based on a sample of U.S. men and women age 16 or older from 1948 to 2010.

trajectories are still concave, but less strongly so in the sense that labor force participation is higher in the twenties than was the case for earlier cohorts.

The central pattern for purposes of this paper is that, comparing across cohorts, labor force participation rose up through the 1952 cohort and stopped rising for subsequent cohorts.[6] This is a very strong pattern for women in their thirties or older. For example, the FLFP rate in age group 35-39 was 60 percent for the 1940 cohort, 68 percent for the 1946 cohort, and 76 percent for the 1952 cohort, and then it essentially has held steady for subsequent cohorts: 76 percent for the 1958 cohort, 77 percent for the 1964 cohort, and 75 percent for the 1970 cohort.

The exception to this pattern is that labor force participation for younger women continued to rise up through the 1970 cohort and then leveled off. Later in this paper, I will examine the roles of delayed marriage and child rearing in accounting for that pattern.

To examine what happened between the intervening cohorts around the plateau, Figure 2.2.B presents the trajectories for the cohorts 1952-1958. The figure confirms that the rise in FLFP for women in their thirties and older was brought to an end by the cohorts born after 1952. For example, for women aged 35-39, the FLFP rates for the 1952, 1953, and 1954 cohorts were 76, 75, and 76 percent, respectively.

2.3.B Educational attainment

As noted by many writers (for example, Goldin, Katz, and Kuziemko, 2006), the rise in women's labor force participation was accompanied by a rise in women's educational attainment. In this section, I reaffirm the trend in women's educational attainment, and I explore its connection to the rise and plateau in women's labor force participation.

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[6] Fallick and Pingle (2007) also point out that FLFP leveled off with the cohorts born in the early 1950s.

I divide the sample into three education categories: high school or less, some college, and college completion or more. Women with high school or less education represent female high school dropouts and graduates; women with some college education include those who had some college experience but no bachelor's degree; women with college completion or more denote women with a bachelor's degree.[7]

Figure 2.3 shows the fraction of women by educational attainment and birth cohort, for age group 25-29. The figure confirms the familiar finding that women's educational attainment has risen dramatically across cohorts. While the high school fraction has declined significantly, both fractions of some college and college completion or more have risen in parallel. For example, 64 percent of women born in 1948 had not gone to college as of their late twenties. Women born 33 years later, in contrast, showed a complete reversal: about 64 percent of women had attended college by the same age. For the 1965 and subsequent cohorts, more than half of women attended college in this age group. In terms of time period, this indicates that by the early 1990s, when the plateau in FLFP appeared, the fraction of women with college education began to exceed the fraction without.

Interestingly, unlike the FLFP rate, the fraction attending college continued to rise beyond the cohorts born in the early 1950s. The fraction for women aged 25-29 continued to rise until cohorts born in the early 1970s, and then it leveled off. An important topic for further research is why the plateau in women's college-going occurred 20 years after the plateau in labor force participation. Lee (2012) suggests that the trend towards higher work expectations of

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[7] In order to construct a measure of educational attainment that is consistent over my sample period, I adopt the methodology proposed by Jaeger (1997). Because of a 1992 change in CPS coding of educational attainment, the trend in high school graduation before 1992 is not comparable to the trend after 1992. I therefore merge the high school dropouts and graduates into one category.

young women across birth cohorts may partly account for the upward trends in women's college attendance and completion.

Figure 2.4 presents more information on the connection between the trends in FLFP and women's educational attainment. The three panels of Figure 2.4 display FLFP rates within each education category. Comparing the three panels shows that FLFP rates regularly have been greater for more educated women. Comparing cohorts within each panel shows that both the rise and plateau in FLFP appear similarly in all education categories. The figure displays the difference between younger age groups across education categories for cohorts born after 1952. The participation rate was higher in the late twenties than in the early thirties for women with college education, while the opposite was true for women without college education. In combination, these comparisons suggest that trends in FLFP must be related to both the within-education-category trends and the changing fraction of women with college education.

To quantify the connection between trends in FLFP and educational attainment, I proceed to a "shift-share" analysis. Let y^d be a dummy variable that equals one if a woman in group d is in the labor force and zero otherwise. For now, d is a multiple-category variable that equals one if a woman has high school education, two if she has college experience, and three if she has a bachelor's degree. (In a later section, d will have categories more than three involving various combinations of college-going, marriage, and child rearing.) Note that the overall FLFP rate $\bar{y}$ can be written as $\sum_d \pi^d \bar{y}^d$, where π^d denotes group d's population share and $\bar{y}^d$ is group d's labor force participation rate. Then the change in the overall FLFP rate between cohorts s and t can be decomposed as:

$$
\begin{aligned}
\bar{y}_t - \bar{y}_s &= \sum_d \pi_t^d \cdot \bar{y}_t{}^d - \sum_d \pi_s^d \cdot \bar{y}_s{}^d \\
&= \sum_d \pi_t^d \cdot \bar{y}_t{}^d + \sum_d \pi_t^d \cdot \bar{y}_s{}^d - \sum_d \pi_t^d \cdot \bar{y}_s{}^d - \sum_d \pi_s^d \cdot \bar{y}_s{}^d \\
&= \underbrace{\sum_d \bar{y}_s{}^d \cdot (\pi_t^d - \pi_s^d)}_{\text{Composition effect}} + \underbrace{\sum_d \pi_t^d \cdot (\bar{y}_t{}^d - \bar{y}_s{}^d)}_{\text{Within-group effect}} \qquad (1)
\end{aligned}
$$

The first element in the final expression indicates the composition effect, which is the change in labor force participation attributable to the change in the population share of each education category. The second component represents the within-group effect, which is the difference attributable to the changes in labor force participation within each education category.[8]

Figure 2.5 summarizes the results of this decomposition for ages 35-39 and ages 25-29 since the decomposition results for women younger and older than 30 diverge substantially. Each decomposition is performed in a way that the FLFP rate of a given cohort is compared to the FLFP rate of a fixed base cohort. The base cohort is the earliest cohort observed in the data for a given age group. Thus, in Figures 5, 8 and 12, the base cohorts for age groups 35-39 and 25-29 are the 1933 and 1943 cohorts, respectively. Each decomposition figure contains four lines. The top solid line shows the FLFP rate by cohort, and the bottom solid line presents the FLFP rate of the base cohort, which is by definition a horizontal line. The broken red line and the solid blue line with a mark represent within-group and compositional changes, respectively. The composition effect for a given cohort plots the composition effect in equation (1) holding the

[8] Of course, there are numerous ways of decomposing. Alternatively, equation (1) can be written as $\bar{y}_t - \bar{y}_s = \sum_d \bar{y}_t{}^d \cdot (\pi_t^d - \pi_s^d) + \sum_d \pi_s^d \cdot (\bar{y}_t{}^d - \bar{y}_s{}^d)$. I have performed decompositions using this alternative equation, and it turns out that the decomposition results are robust to ways of decomposing.

within-group FLFP rates constant at the values for the base cohort. Thus, it shows what the FLFP rate would have been following the base cohort if there were only compositional changes. Similarly, the within-group change plots the within-group effect from equation (1) holding constant the population share of each education category.

The top panel of Figure 2.5 shows that, for ages 35-39, the rise in FLFP up through the cohorts of the early 1950s was driven almost entirely by the within-group effect, that is, by rising FLFP among all those with high school or less, some college, and college or more education. And the plateau for subsequent cohorts reflects the plateaus within all education categories. The bottom panel shows that, for ages 25-29, the rise in FLFP up through the cohorts of the early 1970s was driven predominantly by the within-group effect, but also to some extent by the rising shares of college attendees and graduates. The plateau after the early 1970s reflects leveling off in both the within-group labor force participation rates and the fractions going to college.

To describe the quantitative magnitudes more precisely, I provide Table 2.1 as a supplement to the decomposition figures. I pick up two arbitrary sets of cohorts, the 1946, 1958, and 1970 cohorts for age group 35-39 and the 1956, 1968, and 1980 cohorts for age group 25-29, and report the estimates for the cohorts shown in Figures 5, 8, 11, and 12 in Panels A, B, C, and D of Table 2.1, respectively. Columns (1), (2), and (3) respectively present the change in FLFP between the base cohort, the 1933 cohort, and each cohort from the former set; columns (4), (5), and (6) respectively present the change in FLFP between the base cohort, the 1943 cohort, and each cohort from the latter set. Column (3) in Panel A of Table 2.1 indicates that seven percent (=1.75/26.01) of the total change in FLFP between the 1933 and 1970 cohorts for age group 35-39 was due to the changes in educational composition of women, whereas column (6) in the same panel implies that 16 percent (=5.02/30.64) of the change between the 1943 and 1980 cohorts for age group 25-29 was. Detailed within-group effects by category indicate that the within-group effects reported in both columns came almost evenly from the three education categories.

2.3.C. Marriage

Marriage patterns also have changed dramatically over the past half century (Isen and Stevenson, 2010). Figure 2.6.A shows the fraction of women ever married (currently married, divorced, separated, or widowed) by age and birth year.[9] The fraction plummeted for all age groups. The particularly large declines for women under age 30 reflect a trend towards later marriage. About 65 percent of women born in 1946 were ever married as of age 20-24. For women of the same age but born 30 years later, only 28 percent had been married. Since this sharp decline was accompanied by a rise in the share of college-educated women, one might think that marital delay is associated with women's college-going behavior. That conjecture is partly true because the share of ever-married women has been lower among those that attended college for most cohorts. Figure 2.6.B, however, shows that the share of ever-married women has not dropped more for women with college than for women without college.   For cohorts born in the late 1960s, the share of ever-married women was even higher among those that attend college as of age 35-39. For cohorts born before 1970, the share declined more rapidly among those that complete college as of age 25-29, but the share for each education category showed a similar pattern of decline for the subsequent cohorts.

Figure 2.7 displays labor force participation rates by age and birth cohort disaggregated between ever-married and never-married women. The figure shows that never-married women

[9] The figure does not include the series for women aged 45 or older because it is not much different from the series for women aged 40-44.

traditionally have participated in the labor force more than ever-married women, but that the upward trend in FLFP up through the 1952 cohort was especially pronounced for ever-married women. Again, the two figures in combination suggest that both within-group trends and composition changes must have contributed to changes in FLFP.

Again, I perform a shift-share analysis to quantify the roles of within-group and composition changes. The framework is still as shown in equation (1), but this time with a binary categorization between ever-married and never-married women. The results are summarized in Figure 2.8. The top panel shows that, for women aged 35-39, the upward trend in labor force participation up through the cohorts of the early 1950s was driven almost entirely by rising labor participation within categories, but the detailed within-group effect reported in Panel B of Table 2.1 indicates that it is completely driven by ever-married women.

The lower panel shows that, for women aged 25-29, the upward trend in labor force participation up through the cohorts of the early 1950s was mostly driven by within-group trends, but also reflects the increasing share of women still not married by their late twenties. The continuing rise in FLFP after those cohorts entirely reflects a continued trend away from marriage by the late twenties. Column (6) in Panel B of Table 2.1 indicates that this trend contributed 46 percent (=14.10/30.64) of the change in FLFP between the 1943 and 1980 cohorts for age group 25-29.

2.3.D Child rearing

Women's labor force participation is strongly related to the presence of young children in the household, so I will define "child rearing" as the presence of children under the age of six.[10] Figure 2.9 shows age-specific trends across cohorts in this measure of child rearing. The striking pattern is a trend towards later child rearing: child rearing in the twenties has declined, while child rearing in the thirties has increased. Of course, this pattern is connected to the trend towards later marriage. As women in more recent cohorts marry and have children later than those in earlier cohorts, the contrast across age groups becomes smaller. For example, among women born in 1946, 60 percent had young children while in their late twenties, compared to only 22 percent while in their late thirties. For women born 24 years later, 41 percent had young children while in their late twenties and 32 percent did in their late thirties.

Figure 2.10 shows labor force participation rates by age and cohort, disaggregated between women who do and do not have children under the age of six. A comparison of the two panels confirms that labor force participation is much lower among women with young children. The lower panel shows that, among women without young children, the rise in labor force participation leveled off after the 1952 cohort. In contrast, the upper panel shows that, among women with young children, labor force participation continued to rise at least through the cohorts of the early 1970s.

Figure 2.11 repeats the shift-share analysis, this time with a categorization between women with and without young children. The lower panel shows that, for women aged 25-29, the rise in labor force participation up through the cohorts of the early 1970s was driven mostly

[10] Although the age of six is a conventional cutoff, it is admittedly arbitrary. However, results using other age cutoffs (the ages of two, three, four, five and seven) do not change much from those in Figures 8 and 9.

by upward within-group trends, but also by an increasing fraction of women without young children. The point can be more clearly illustrated using estimates shown in Panel C of Table 2.1. The estimates in column (5) in Panel C imply that 72 percent (=21.97/30.42) of the change in FLFP between the 1943 and 1968 cohorts was driven by the within-group effect, with a bigger within-group effect from women with young children. The other 28 percent of the change was driven by the change in women's child-rearing composition between the two cohorts.

The upper panel in Figure 2.11 tells a different story about women aged 35-39. For them, the rise in labor force participation up through the cohorts of the early 1950s was driven almost entirely by rising labor participation within both groups, women with and without young children. For subsequent cohorts, the within-group effect actually *exceeds* the overall change in labor force participation. That is, the labor force participation rate would have been higher if not for a negative composition effect due to the increasing share of women with young children in this age group. The estimates in column (1) indicate that 87 percent (=16.51/19.01) of the change in FLFP between the 1933 and 1946 cohorts was driven by the within-group effect. The estimate of the composition effect in column (3) implies that the FLFP rate would have fallen between the 1933 and 1970 cohorts by 0.4 percentage points, had it not been for the within-group effect.

2.3.E The joint contribution of education, marriage, and child rearing

Of course, the separate shift-share analyses in the previous sections overlook the possibility of interactions among the roles of education, marriage, and child rearing. This section extends the decomposition in equation (1) by using twelve categories: every combination of the three education categories, ever-married or never-married, and with or without children under the age of six.

The results, shown in Figure 2.12, are much as one would guess from the separate decompositions in the previous sections. For women aged 35-39, the rise in labor force participation up through the cohorts of the early 1950s is driven almost entirely by rising labor force participation within categories. And the plateau for subsequent cohorts reflects a leveling off of labor force participation within categories. Column (3) in Panel D of Table 2.1 indicates that 14 percent (=3.65/26.01) of the change in FLFP between the 1933 and 1970 cohorts was due to the composition effect. Decomposition of the within-group effect for this age group, although not reported in the table, shows that 74 percent of the within-group change in FLFP between the 1933 and 1946 cohorts was driven by married women without young children and 11 percent of the within-group change was driven by college-educated married women with young children. The upward within-group trend among the latter group of women is quantitatively more important in explaining the change in FLFP between the 1933 and 1970 cohorts, with 32 percent of the change driven by the latter group.

For women aged 25-29, the rise in labor force participation up through the cohorts of the early 1950s is driven by both within-group trends and by composition effects, with the latter mainly associated with the trends away from early marriage and childbearing. The subsequent rise in labor force participation up through the cohorts of the early 1970s is mostly driven by the continuation of the trends away from early marriage and childbearing. Column (6) in Panel D of Table 2.1 indicates that 46 percent (=14.21/30.64) of the change in FLFP between the 1943 and 1980 cohorts was driven by the composition effect. Decomposition of the within-group effect for this age group suggests that within-group change among high-school-educated married women with young children is quantitatively the most important in explaining the within-group change in FLFP between the 1943 and 1956 cohorts, with 39 percent of the within-group change driven by this group. The rate drops to 20 percent for the within-group change in FLFP between the 1943 and 1980 cohorts.

2.4 Summary and discussion

After many decades of trending upward, the female labor force participation rate in the United States hit a plateau in the early 1990s. Using data from March Current Population Surveys, I have documented salient aspects of this important social change. I have shown that, to a first approximation, the plateau can be parsimoniously characterized as a leveling off of labor force participation for birth cohorts from the early 1950s on. My series of shift-share analyses shows that, both the plateau and the earlier upward trend appeared within every category broken down by education, child rearing, and marital status except that the upward within-group trend did not appear among never-married women.

The main qualification to that simple summary is that, for women under the age of 30, labor force participation continued to rise up through the cohorts of the early 1970s. The shift-share analyses show that the continued rise for younger women was intertwined with the trends towards later marriage and childbearing. I also have shown that the trend towards more college education among women continued beyond the cohorts of the early 1950s. The composition effects from educational attainment, marriage, and child rearing respectively contributed 7, 11, and -2 percent of the changes in FLFP between the 1933 and 1970 cohorts for age group 35-39. The corresponding contributions to the change in FLFP between the 1943 and 1980 cohorts for age group 25-29 are 16, 46, and 33 percent.

Of course, all of these findings beg the question of *why* these interrelated trends have developed in the ways that they have. Goldin (2006), Juhn and Potter (2006), and Boushey (2005) speculated that the plateau in women's labor force participation may have been a temporary response to the recession of the early 2000s, but my updating of the evidence refutes that conjecture. To my knowledge, the most ambitious attempt to identify causes of the plateau is the study by Fortin (2009), which argues that women in more recent cohorts have reverted to more traditional gender role attitudes than women in preceding cohorts and that this change in gender role attitudes explains at least a third of the plateau in FLFP.

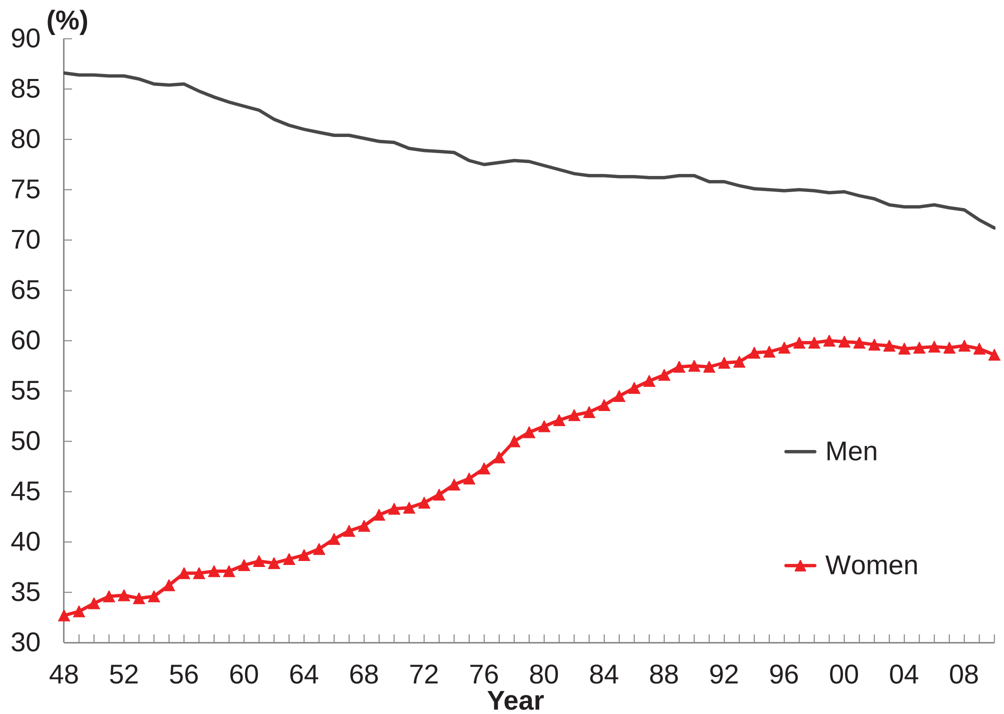
Clearly, more research is needed. My hope is that, by providing a transparent summary of the salient facts about recent developments in women's labor force participation and related socioeconomic outcomes, the present paper will stimulate and guide further research that will delve more deeply into the causes of the shifting trends. For example, if real wage growth and advances in home production technology are credited with explaining the long-running rise in women's labor force participation, why did those forces become inoperative for cohorts born after the early 1950s?

Table 2.1. Decomposition in difference in FLFP rates between birth cohorts using socioeconomic outcomes of women

Age group	35-39			25-29		
	Cohorts between 1933 and			Cohorts between 1943 and		
	1946	1958	1970	1956	1968	1980
	(1)	(2)	(3)	(4)	(5)	(6)
A. Decomposition using educational attainment						
Total change in FLFP	19.01	26.79	26.01	25.32	30.42	30.64
Composition effect	0.65	0.65	1.75	1.87	3.53	5.02
Within-group effect	18.35	26.15	24.26	23.44	26.88	25.62
High school or less	10.08	10.15	7.11	12.73	9.93	8.88
Some college	4.69	10.30	8.53	6.54	10.18	9.67
College completion or more	3.59	5.69	8.63	4.17	6.77	7.06
B. Decomposition using marital status						
Total change in FLFP	19.01	26.79	26.01	25.32	30.42	30.64
Composition effect	0.25	1.76	2.76	5.47	9.84	14.10
Within-group effect	18.75	25.03	23.25	19.84	20.57	16.54
Ever-married	18.44	24.49	22.30	19.46	21.13	16.64
Never-married	0.31	0.55	0.95	0.39	-0.56	-0.10
C. Decomposition using child-rearing status						
Total change in FLFP	19.01	26.79	26.01	25.32	30.42	30.64
Composition effect	2.50	0.66	-0.40	6.98	8.45	10.02
Within-group effect	16.51	26.13	26.41	18.34	21.97	20.62
With children under 6	4.58	10.40	12.09	11.40	13.75	13.33
Without children under 6	11.92	15.73	14.32	6.93	8.22	7.29
D. Decomposition using joint variable						
Total change in FLFP	19.01	26.79	26.01	25.32	30.42	30.64
Composition effect	3.06	2.55	3.65	8.63	11.40	14.21
Within-group effect	15.95	24.24	22.36	16.69	19.02	16.43

Note: The table reports selective estimates shown in Figures 5, 8, 11 and 12. The unit of the estimates is percentage point.

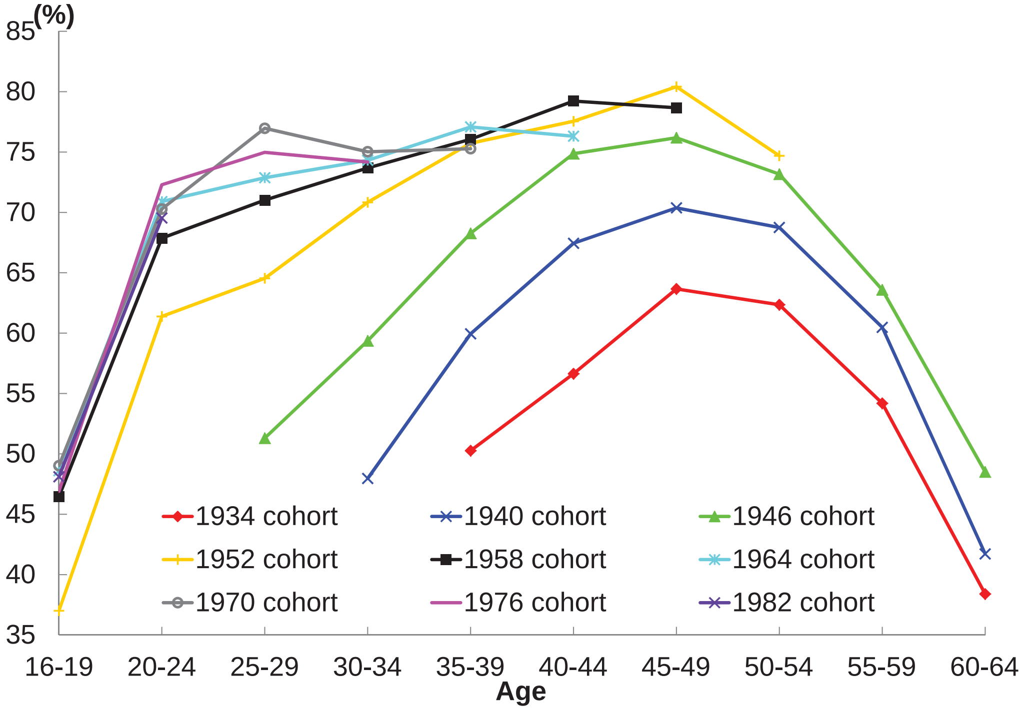
Figure 2.1: Labor force participation rates for men and women



Note: The figure shows civilian labor force participation rates of U.S. men and women age 16 or older. For interpretation of the references to color in this and all other figures, the reader is referred to the electronic version of this dissertation.
Source: BLS series report available at http://data.bls.gov/cgi-bin/srgate

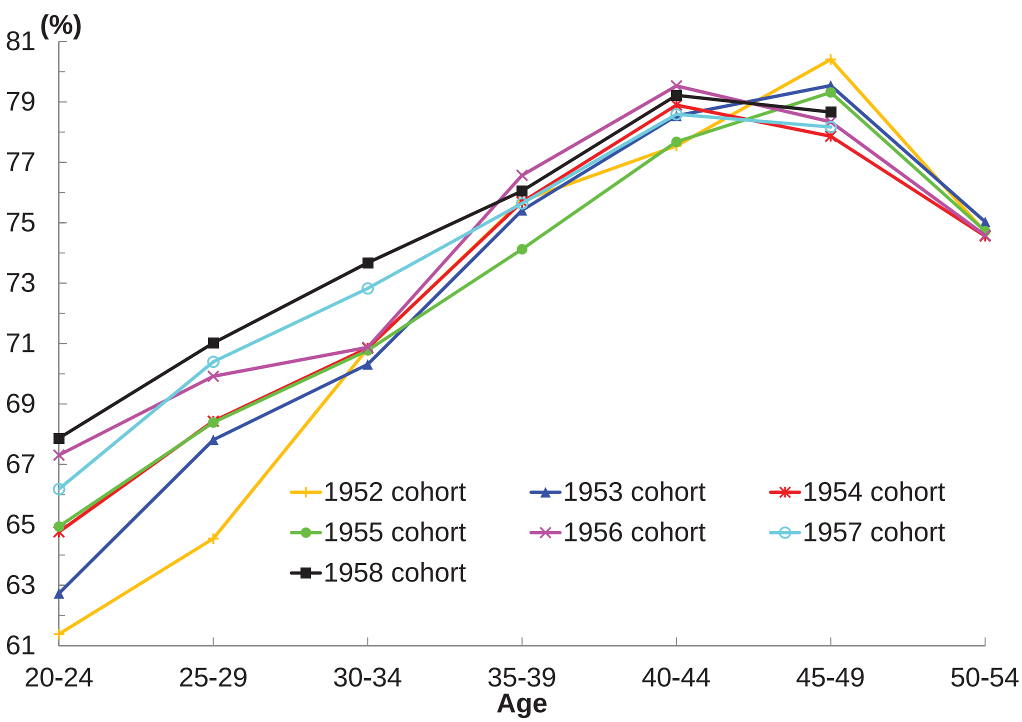
Figure 2.2: Female labor force participation rates by age and cohort

A. 1934 - 1982 cohorts



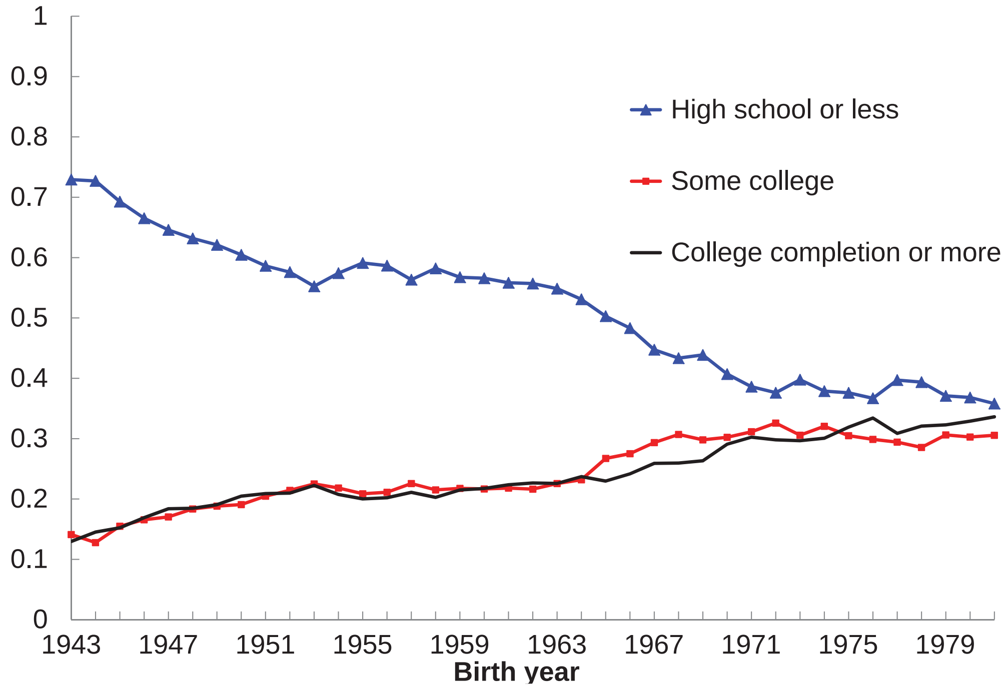
Note: All figures throughout this paper use five-year age bins and single-year birth cohorts.

B. 1952 - 1958



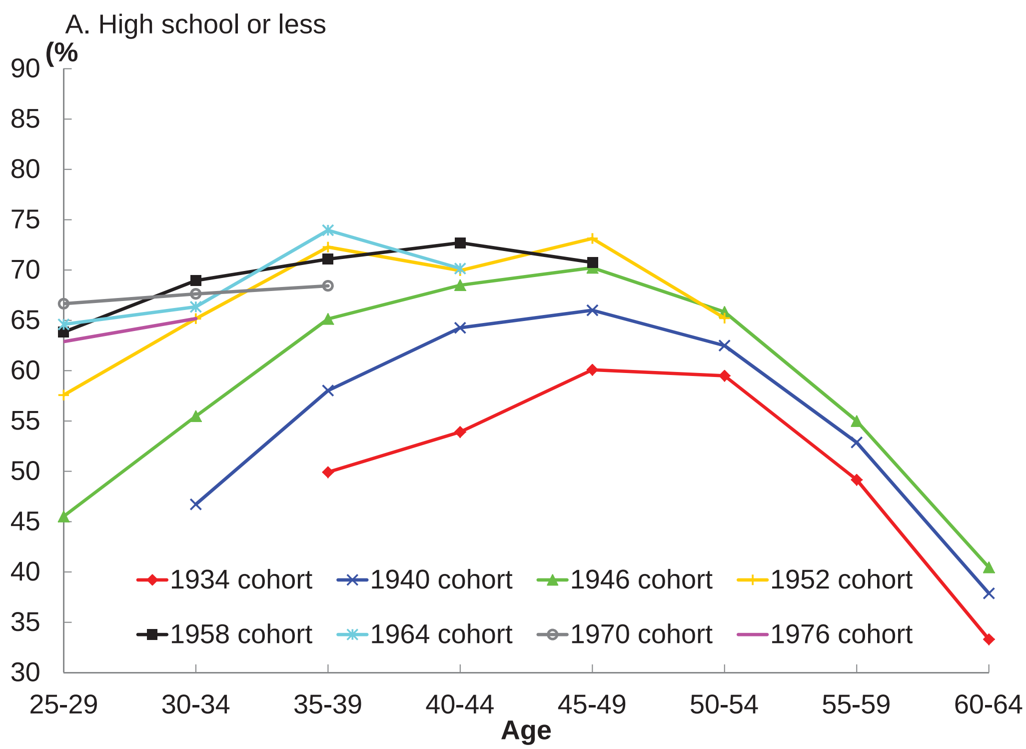
Source: 1968-2010 IPUMS-CPS

Figure 2.3: Women's educational attainment by cohort, ages 25-29



Note: High school or less category includes female high school dropouts and graduates as well as GED recipients. Some college includes women with three or less years of college education and/or women with an associate degree. College completion or more includes women with four or more years of college education and/or women with a bachelor's degree.
Source: 1968-2010 IPUMS-CPS

Figure 2.4: Female labor force participation rates by educational



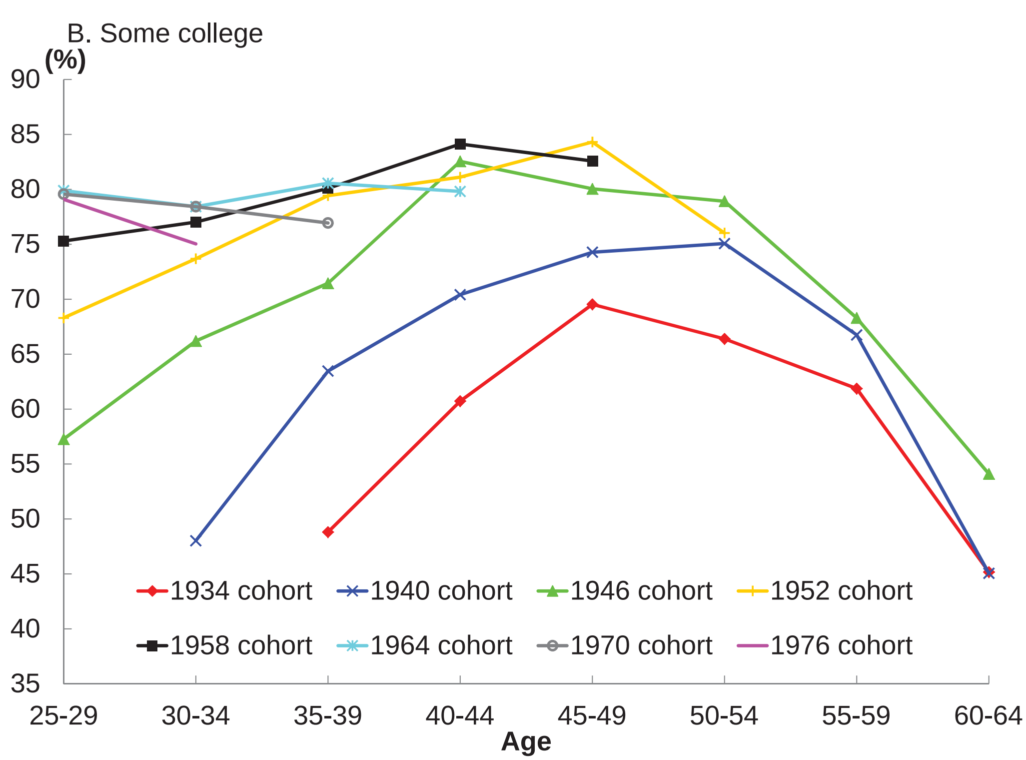
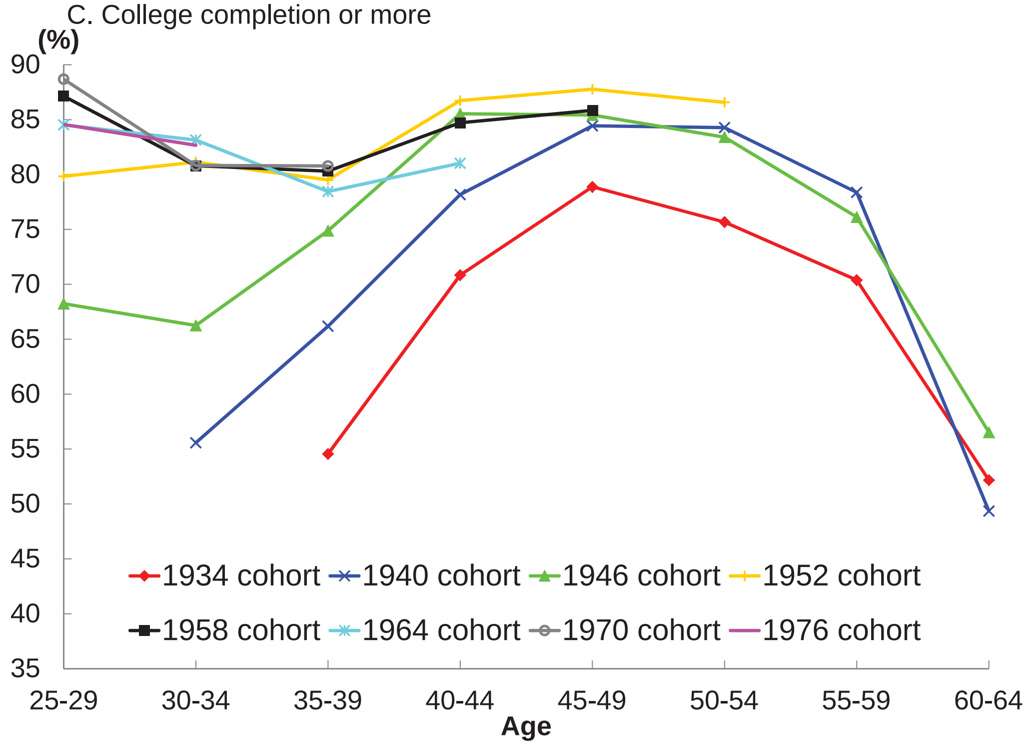


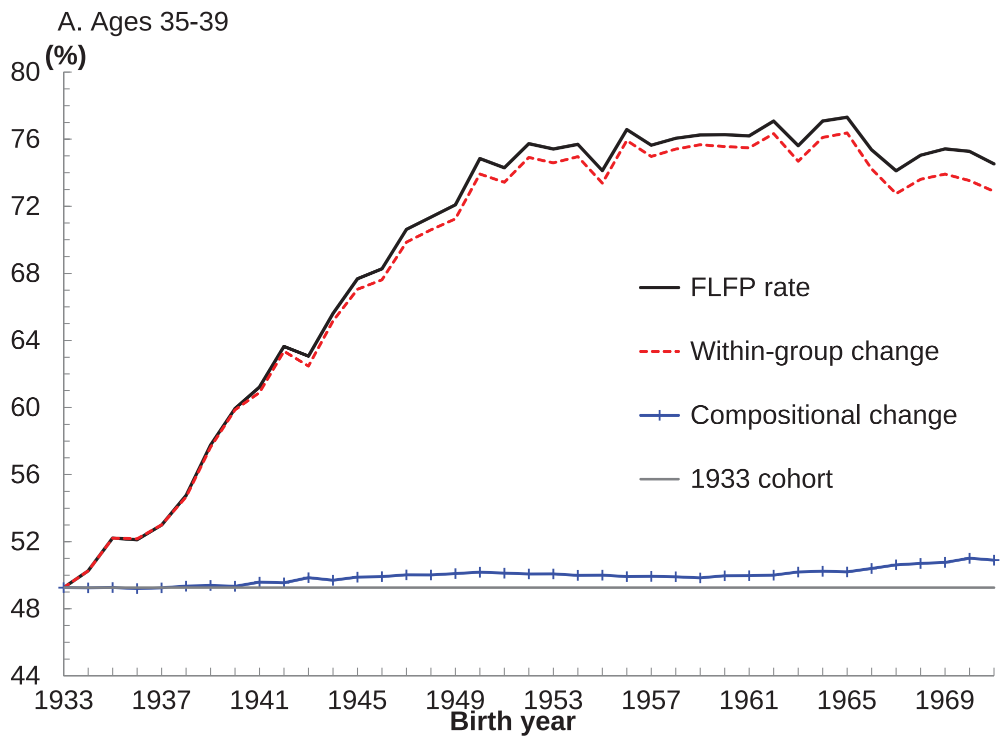
Figure 2.4 (cont'd)

C. College completion or more

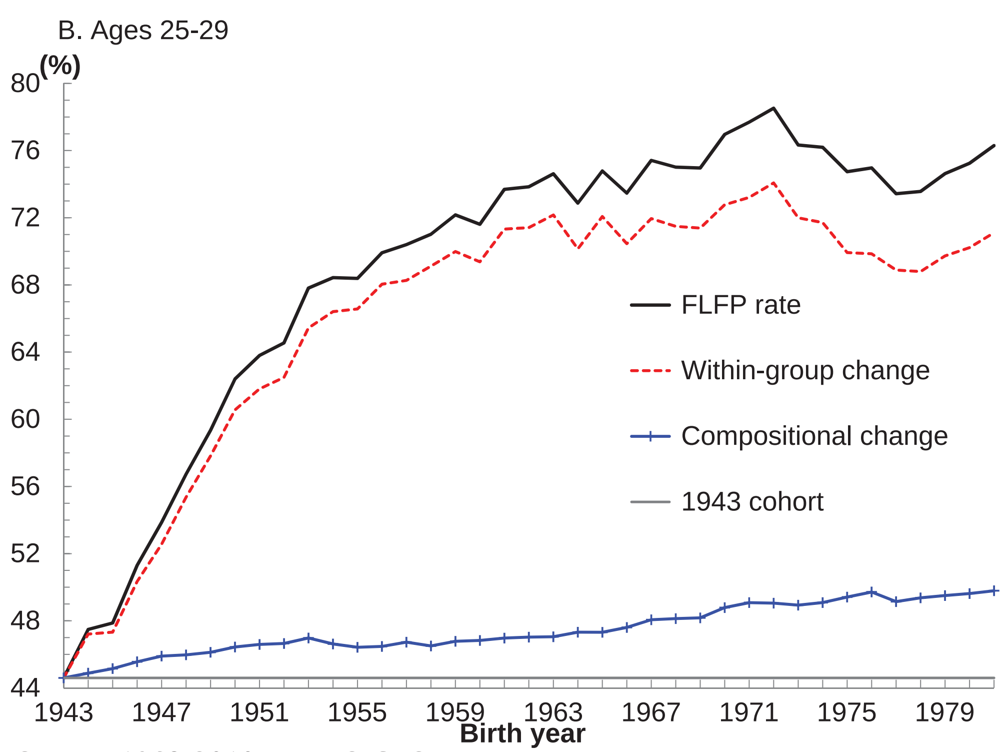


Source: 1968-2010 IPUMS-CPS

Figure 2.5: Decomposition of cohort difference in FLFP using education

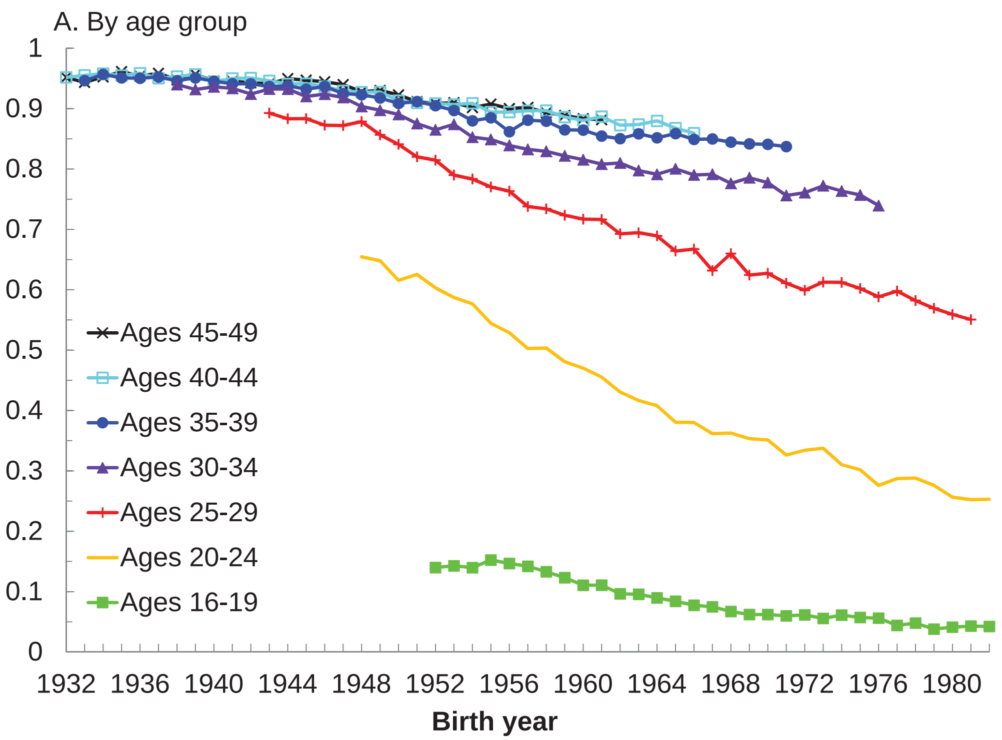


Note: Decomposition is implemented on the FLFP change between base and given cohorts using a multiple-category education indicator. The base cohorts for age groups 35-39 and 25-29 are the 1933 and 1943 cohort, respectively.



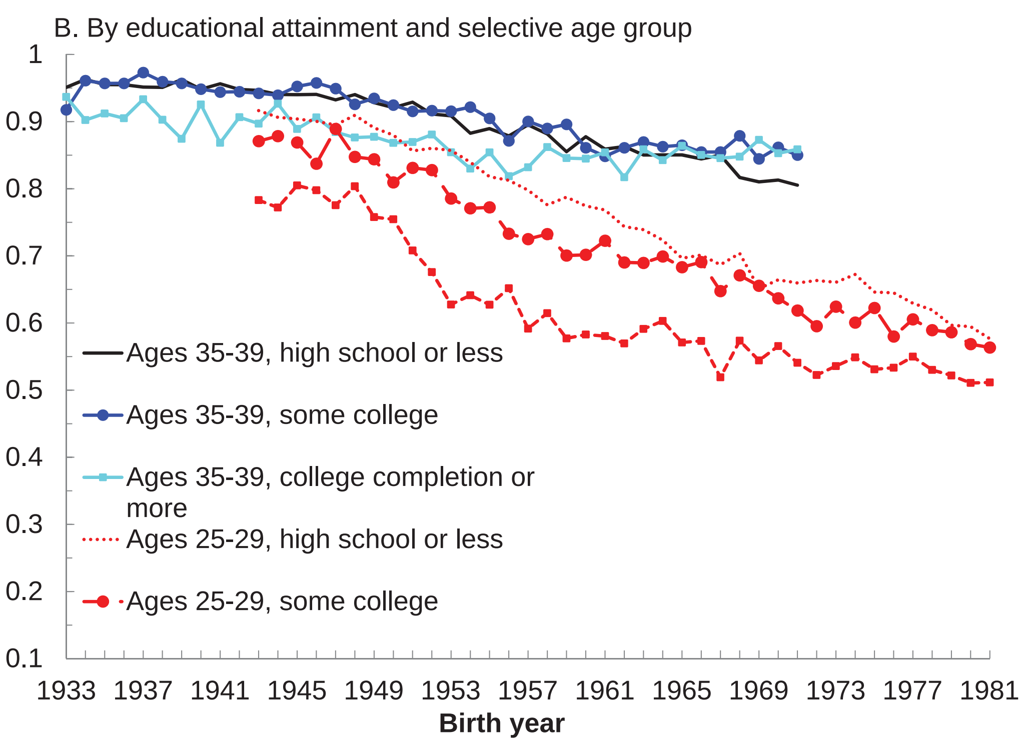
Source: 1968-2010 IPUMS-CPS

Figure 2.6: Fraction of women ever-married



Note: Ever-married women include currently married, divorced, separated and widowed women.

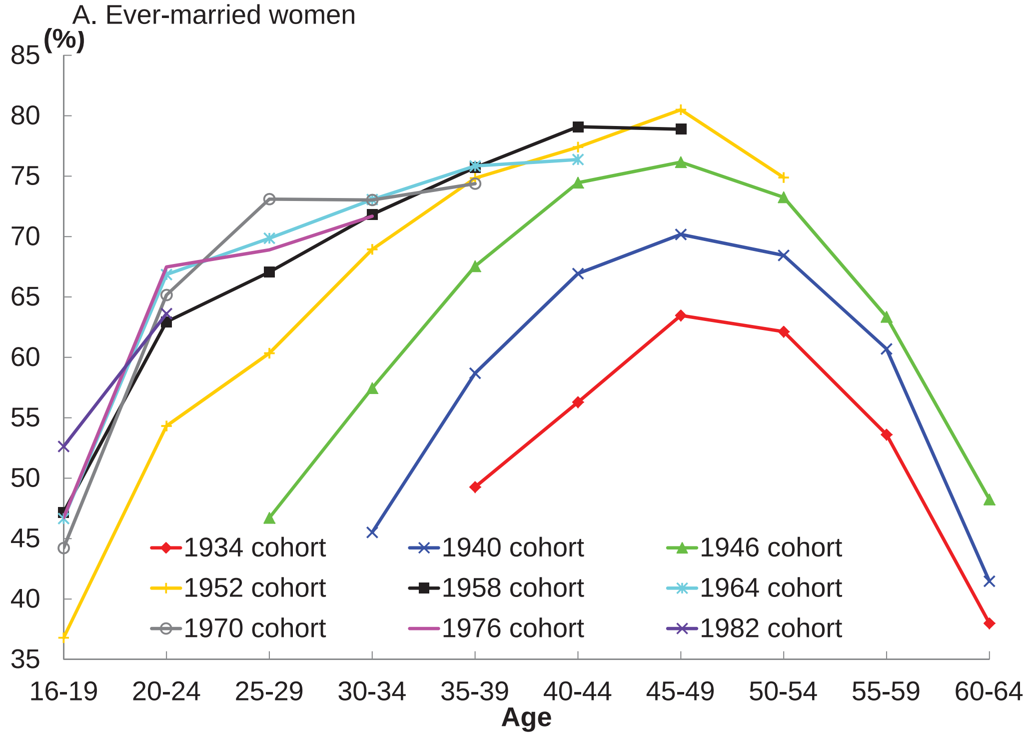
Source: 1968-2010 IPUMS-CPS



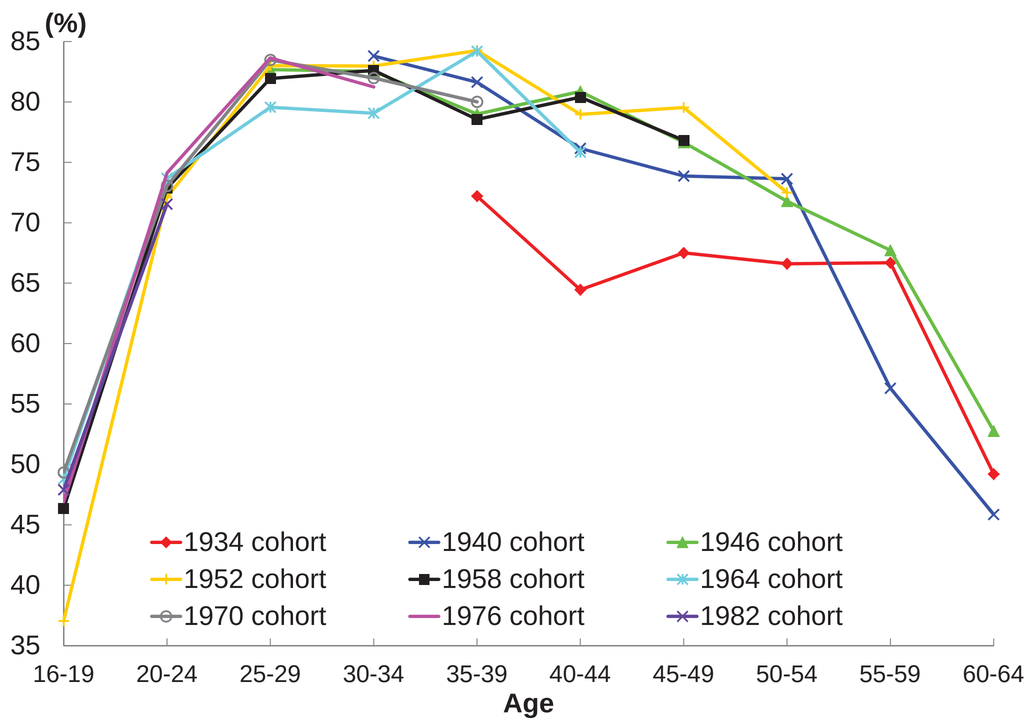
Note: Women with high school include both high school dropouts and graduates. Women with college completion or more mean women with a bachelor's degree.

Figure 2.7: Female labor force participation rates by marital status

A. Ever-married women

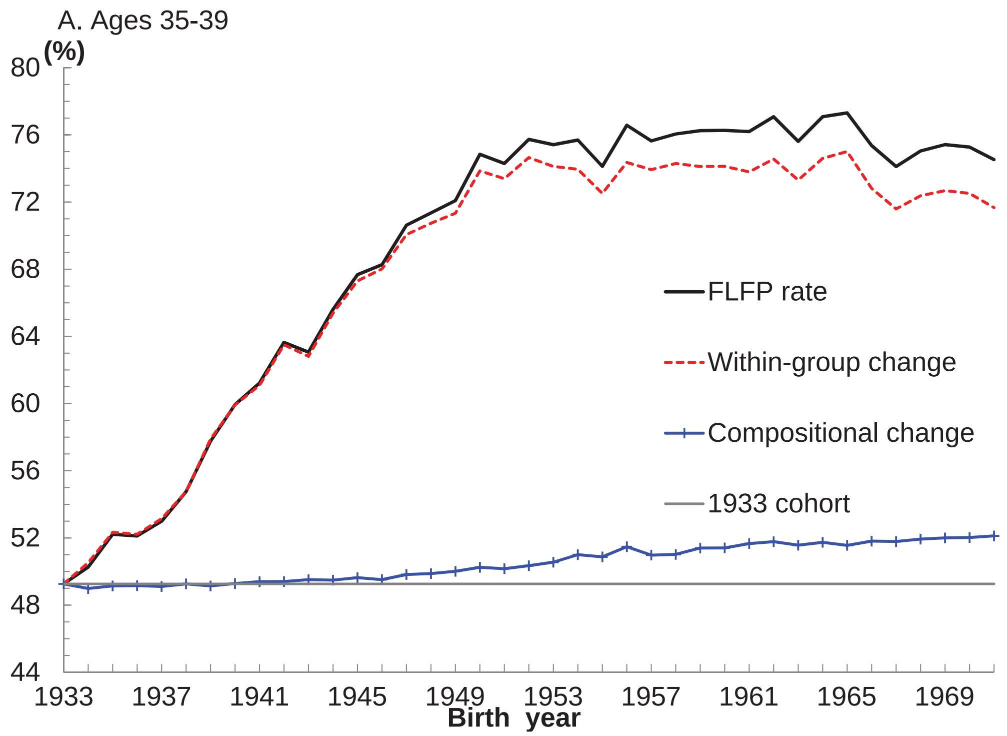


B. Never-married women

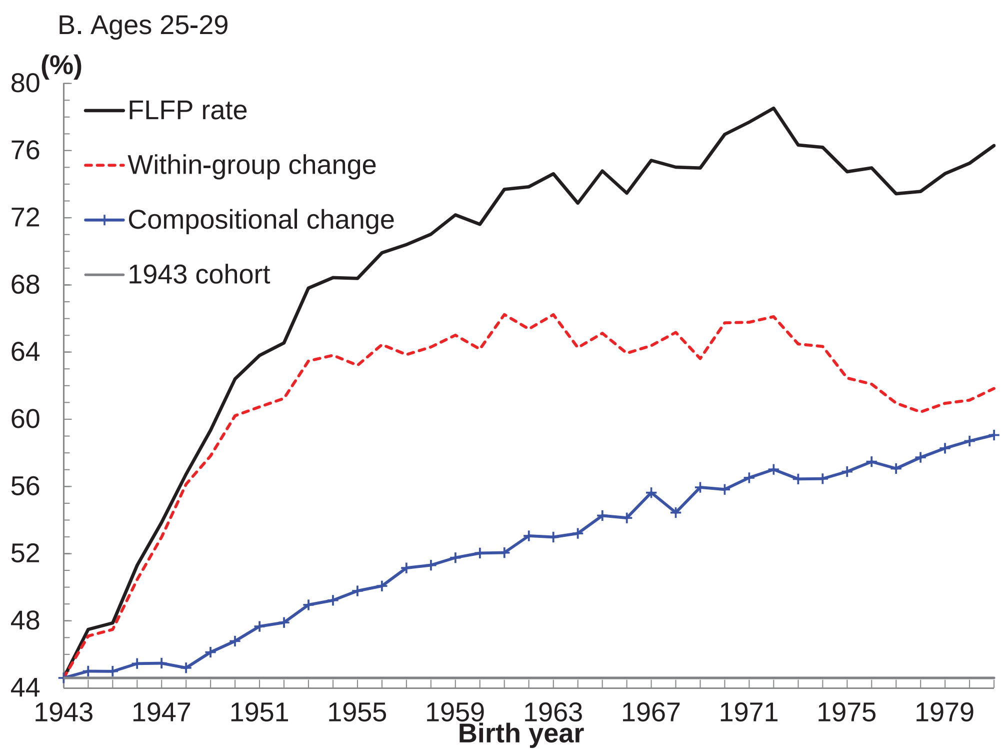


Source: 1968-2010 IPUMS-CPS

Figure 2.8: Decomposition of cohort difference in FLFP using marriage

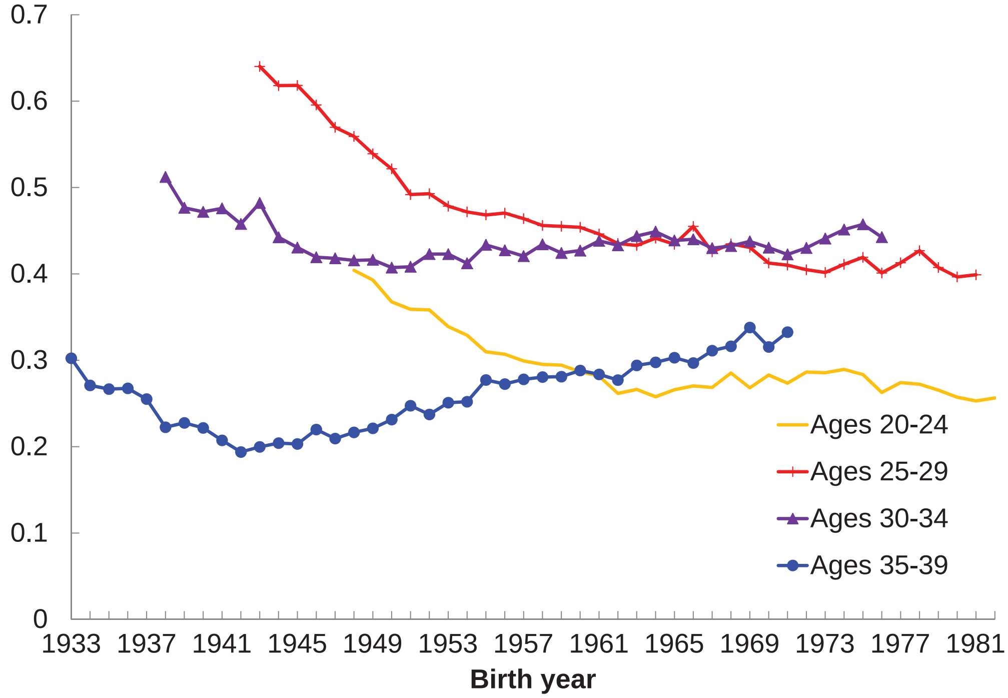


Note: Decompositions are implemented on the FLFP change between base and given cohorts using a dummy variable indicating whether a woman has been ever-married. The base cohorts for age groups 35-39 and 25-29 are the 1933 and 1943 cohort, respectively.



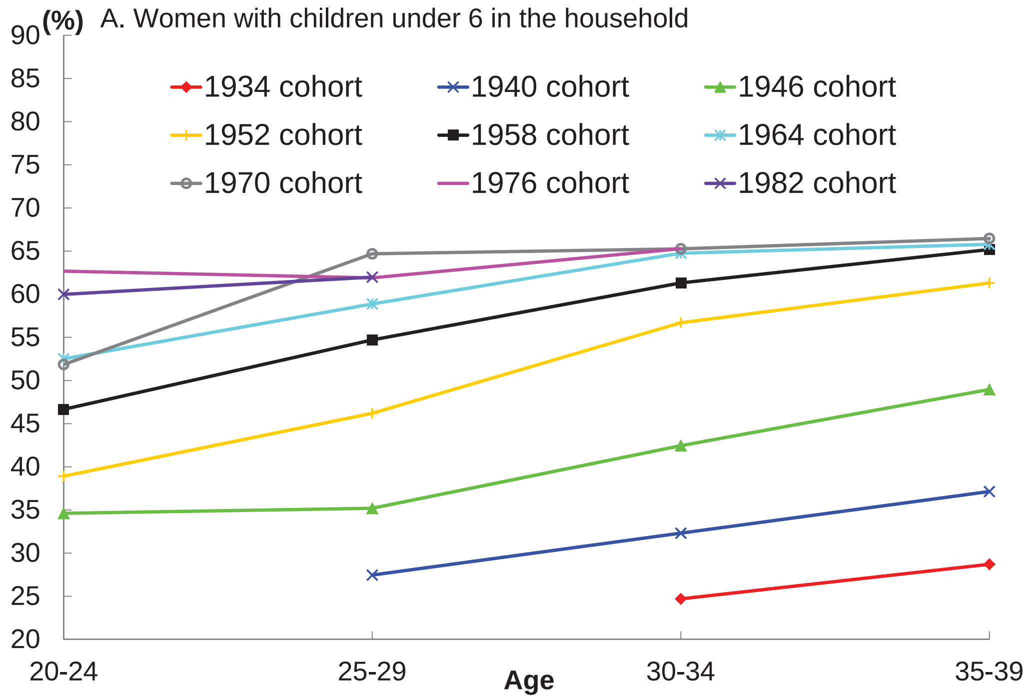
Source: 1968-2010 IPUMS-CPS

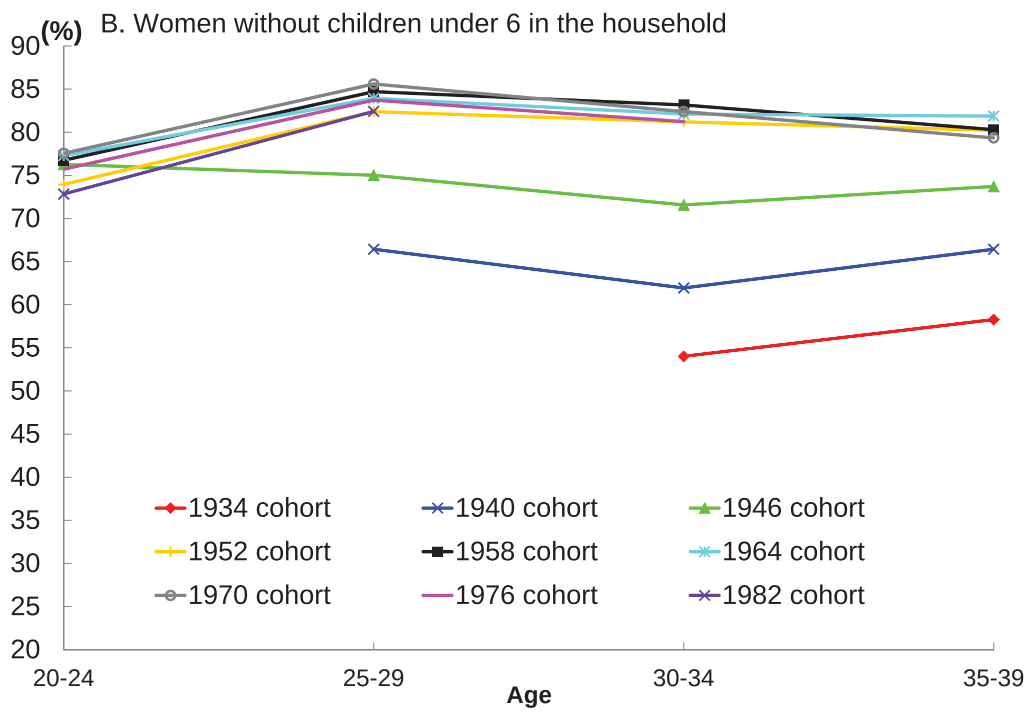
Figure 2.9: Fraction of women who have children under 6 in the household



Source: 1968-2010 IPUMS-CPS

Figure 2.10: Labor force participation rates for women with and without young children

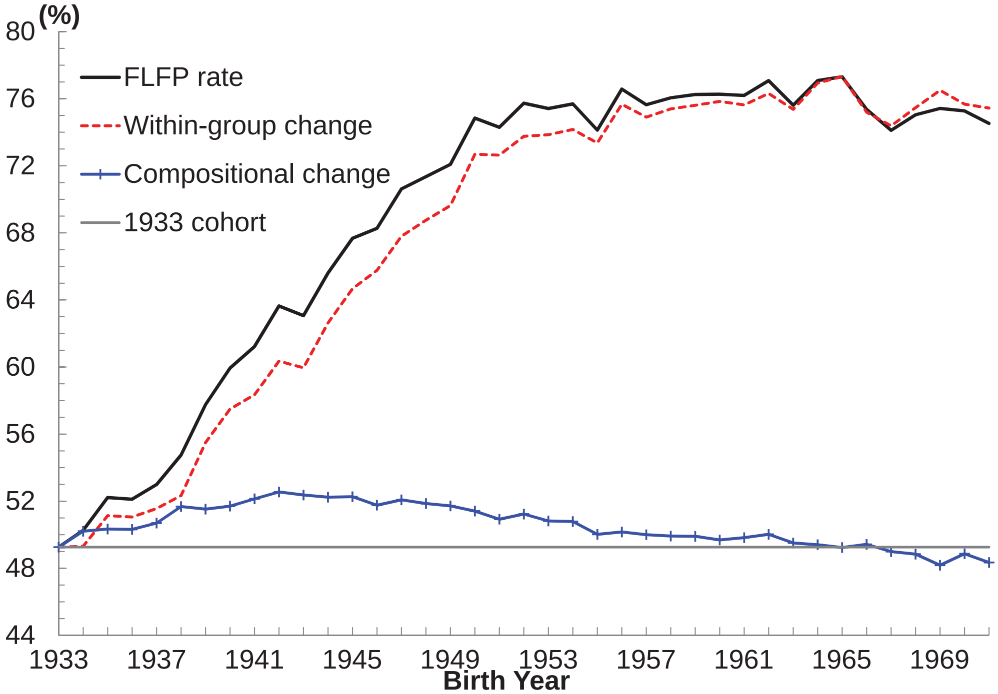




Source: 1968-2010 IPUMS-CPS

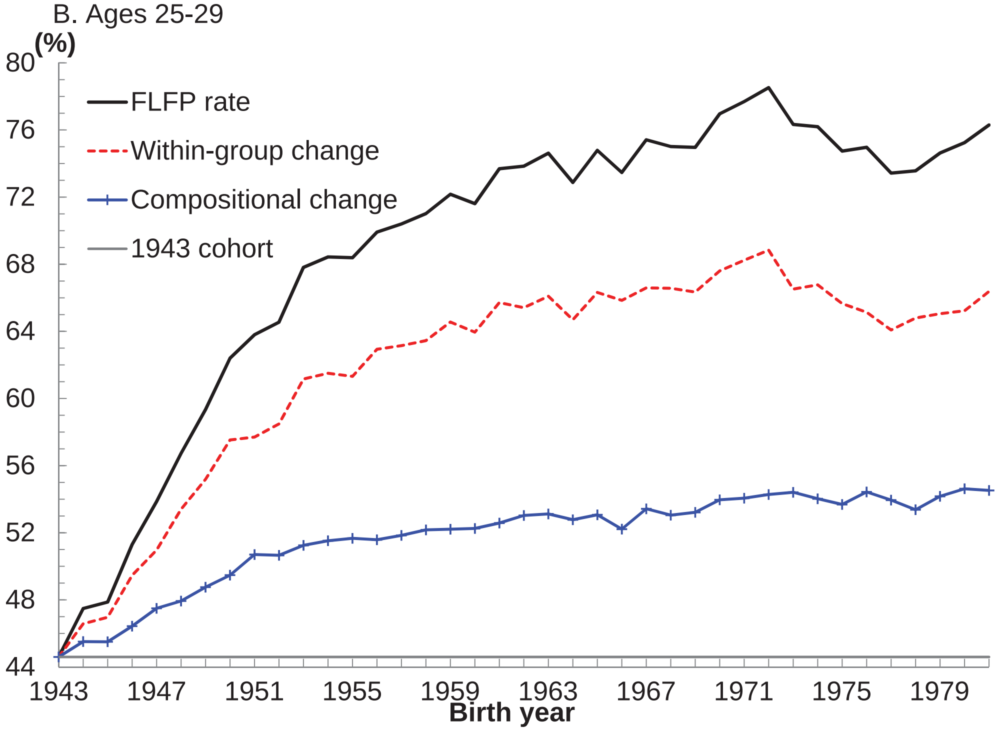
Figure 2.11: Decomposition of cohort difference in FLFP using child-rearing

A. Ages 35-39



Note: Decompositions are implemented on the FLFP change between base and given cohorts using a dummy variable indicating whether a woman has children under 6 in the household. The base cohorts for age groups 35-39 and 25-29 are the 1933 and 1943 cohort, respectively.

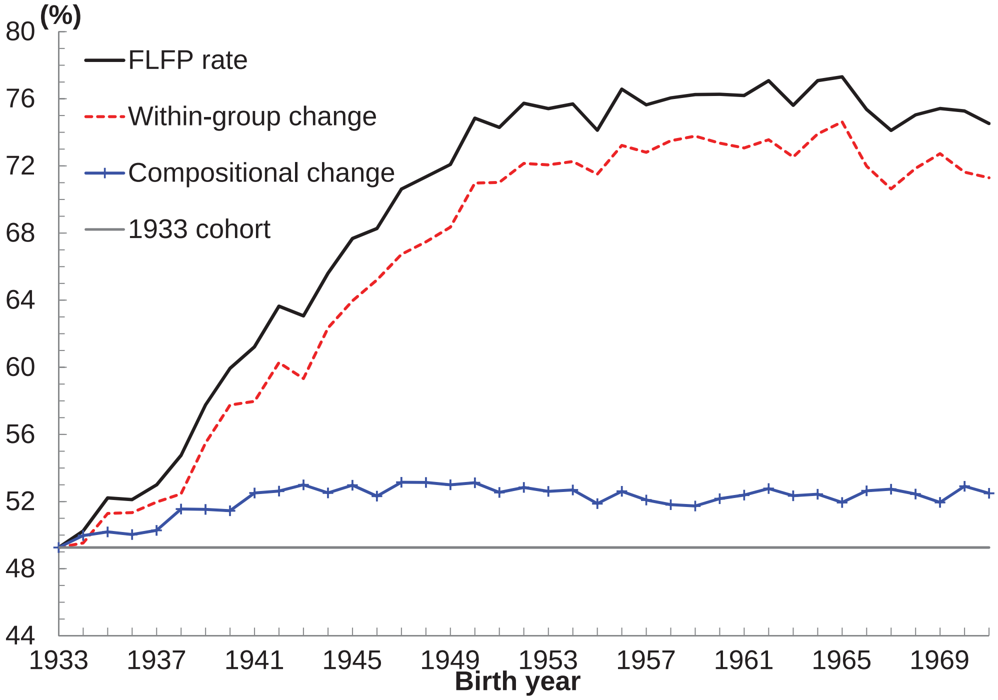
B. Ages 25-29



Source: 1968-2010 IPUMS-CPS

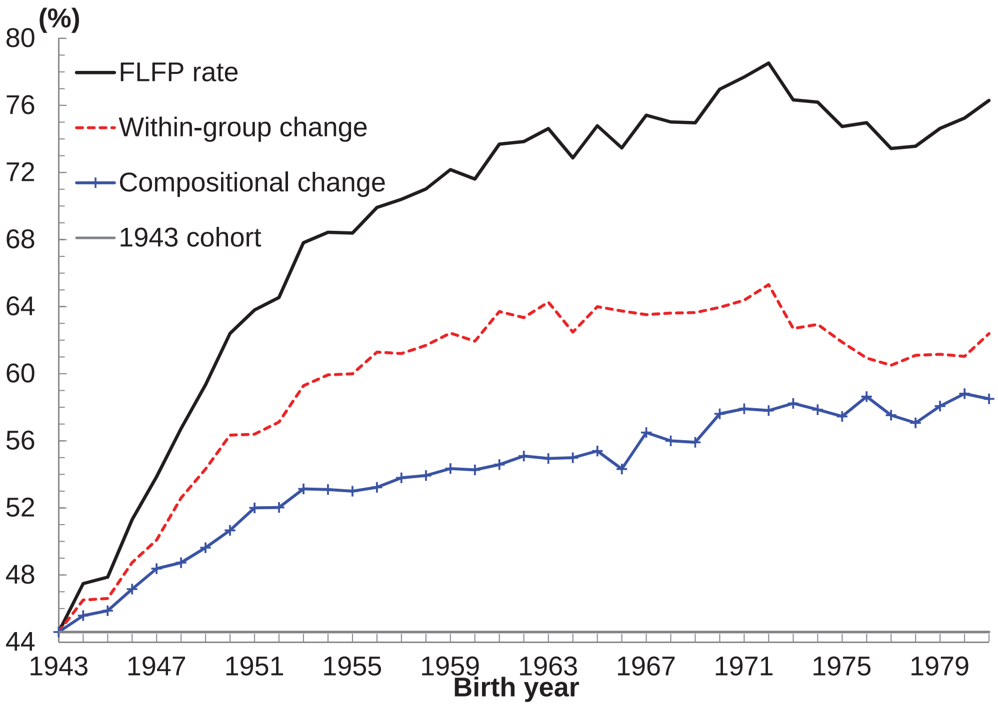
Figure 2.12: Decomposition of cohort difference in FLFP using joint variable

A. Ages 35-39



Note: Decompositions are implemented on the FLFP change between base and given cohorts using a variable which indicates women's educational, marital, and child-rearing status jointly. The base cohorts for age groups 35-39 and 25-29 are the 1933 and 1943 cohort, respectively.

B. Ages 25-29



Source: 1968-2010 IPUMS-CPS

CHAPTER 3

THE UPWARD TREND IN WOMEN'S COLLEGE-GOING: THE ROLE OF TEENAGERS' ANTICIPATED FUTURE LABOR FORCE ATTACHMENT

3.1 Introduction

In recent decades, U.S. women's rates of college attendance and completion have increased so rapidly that women have surpassed men in both rates.[1] As Figure 3.1 shows, both the attendance and completion rates increasingly favor women for birth cohorts from the early 1960s on. Although many studies have suggested reasons that women's college-going exceeds men's, most of those studies are surprisingly unhelpful for explaining why the gender gap has *changed over time* in the way that it has. In this paper, I focus on the *trend* in women's college-going and show that changes in teenagers' anticipated future labor force participation might explain a substantial part of the upward trend in women's college-going.

For example, one strand of the literature on the gender gap in collegiate attainment suggests that non-pecuniary costs of college may be lower for women than men. Becker, Hubbard, and Murphy (2010) and Goldin, Katz, and Kuziemko (2006) argue that greater average non-cognitive skills of women allow women to prepare better for college and to attend and complete college with lower costs. High school graduation rates by gender also provide

[1] The definitions of college attendance and completion used in the Current Population Surveys are slightly different from those used in the National Longitudinal Surveys. For samples from the Current Population Surveys, I define college attendance as at least one year of college education and college completion as four or more years of college education. For samples from the National Longitudinal Surveys, I define college attendance as any college enrollment and college completion as attainment of a bachelor's degree.

evidence of women's better preparedness for college. Heckman and LaFontaine (2010) argue that a decline in the high school graduation rate for men across birth cohorts, in contrast to the stable trend for women, may explain part of the growing gender difference in college attendance.

However, the lower college cost (reflecting higher non-cognitive skills) for women cannot explain the *trend* in college-going by gender because it is not a new phenomenon. As Buchmann and DiPrete (2006) and others note, girls have outperformed boys in school for a long time. Moreover, the high school graduation rate for women, which could be another indicator of non-cognitive skills, has remained stable over the past 40 years (Heckman and LaFontaine, 2010).

Another strand of the literature has suggested that women's college-going exceeds men's because women face a higher wage premium for college education (Charles and Luoh, 2003; Chiappori, Iyigun, and Weiss, 2009; DiPrete and Buchmann, 2006; Dougherty, 2005). Again, however, this observation does not explain the *trend* in college-going by gender because these studies do not indicate that women's wage premium for college has risen faster than men's. In any case, a recent study by Hubbard (2011) casts doubt on the argument that women's wage premium exceeds men's by showing that the apparent difference may be merely an artifact of top-coding bias.

So what *can* explain the trend in women's college-going relative to men's? The closest precedent for the present study is an article by Goldin, Katz and Kuziemko (2006), which argues that changes in young women's anticipated labor force participation may be an important factor in the rise of women's college-going. Using the National Longitudinal Survey of Young Women 1968 (NLSYW68) and the National Longitudinal Survey of Youth 1979 (NLSY79), Goldin et al. find that young women's anticipated participation rose steadily from 1968 to the early 1980s. In addition, they compare college attendance and completion rates between women with future work plans and women without and find that both rates are significantly higher for women with work plans. In this paper, I present more up-to-date trends in young women's work expectations using three National Longitudinal Surveys – the NLSYW68, the NLSY79, and the National Longitudinal Survey of Youth 1997 (NLSY97). In addition, I further develop Goldin, Katz and Kuziemko's insight by providing a simple theoretical model to illustrate the potential importance of young women's work expectations and by presenting a detailed set of regression analyses to test the main implication of the model.

The next section of this paper presents the illustrative theoretical model and explains its practical relevance and testable implications. Section 3.3 describes the NLS data and uses them to document that teenage women's expectations of future labor force attachment have indeed risen across cohorts. Section 3.4 goes on to test empirically whether teenage women's work expectations are positively associated with college-going and uses the results to examine the extent to which rising work expectations can account for the upward trend in women's college-going. Section 3.5 summarizes the findings.

3.2 The potential importance of work expectations for college-going decisions

As noted in the previous section, many studies of gender differences in college-going stress the potential role of the wage rate premium associated with a college education. Most of those studies, however, overlook the interaction of the wage premium with anticipated work hours. To put the point in an extreme way, a high hourly wage premium for a college education would provide little incentive for college-going to a person who expected to work only one nanosecond after college. In this extreme case, even with a high hourly wage premium, the earnings payoff to college would be infinitesimal. The more general point that the expected earnings payoff to college depends on anticipated work hours as well as the hourly wage premium is illustrated by the following simple model.

Suppose woman i's hourly wage W_i is determined by

$$(1) \qquad W_i = e^{\alpha + \beta D_i + \varepsilon_i},$$

where D_i is equal to one if woman i goes to college for four years and zero if she does not go to college. The heterogeneity term ε_i, which reflects idiosyncratic ability, motivation, etc., is observed by woman i, but not by the econometrician. The parameter β is the wage rate premium for college education. Assume that individuals can borrow at interest rate r and that the wage rate premium for college education is large enough so that some women are willing to attend college ($\beta > 4r$).

When deciding whether to attend college or not, woman i compares the expected present discounted value ($E[\text{PDV}]$) with and without college education:

$$(2) \qquad E[\text{PDV with college}] = \int_0^\infty E[H_i] \cdot e^{\alpha+\beta+\varepsilon_i} e^{-rt} dt - T_i$$
$$= (\frac{1}{r}) E[H_i] \cdot e^{\alpha+\beta-4r+\varepsilon_i} - T_i, \text{ and}$$

$$(3) \qquad E[\text{PDV without college}] = \int_0^\infty E[H_i] \cdot e^{\alpha+\varepsilon_i} e^{-rt} dt$$
$$= (\frac{1}{r}) E[H_i] \cdot e^{\alpha+\varepsilon_i},$$

where $E[H_i]$ is woman i's expected hours of work; T_i is her expected PDV of college costs besides foregone earnings, such as tuition and net non-pecuniary costs of college; and lifetimes are long enough that the PDV calculations are approximately correct with an assumption of infinite lifetimes. Then, individual i will choose to attend college if and only if $(2) > (3)$, or equivalently if and only if

$$\text{(4)} \qquad \varepsilon_i > -\alpha + \ln T_i + \ln r - \ln(e^{\beta - 4r} - 1) - \ln E[H_i].$$

Thus, the probability that woman i attends college is

$$\text{(5)} \qquad \Pr[\varepsilon_i > -\alpha + \ln T_i + \ln r - \ln(e^{\beta - 4r} - 1) - \ln E[H_i]] \\ = 1 - F(-\alpha + \ln T_i + \ln r - \ln(e^{\beta - 4r} - 1) - \ln E[H_i]),$$

where $F(\cdot)$ is the cumulative distribution function of ε_i.

This result confirms the conjecture that the probability of going to college depends positively on not only the hourly wage premium for college, but also on the individual's expected future hours of work. Of course, the model is exceedingly simple in many respects. To name just a few examples, it abstracts from the possibility of starting college but not finishing; it ignores the life-cycle evolution of wages and hours after the completion of education; it ignores uncertainty about future wages; and it assumes perfect capital markets. There is no reason to think, however, that incorporating any or all such complications would undo the result that optimizing decisions about college-going should depend on anticipations of future work hours.

It follows that one possible reason for women's increased college-going is a trend towards greater expectations of future labor force attachment among more recent cohorts of young women. At first blush, it seems highly plausible that teenage women's anticipations of future labor force attachment have risen because, as is well known, women's labor force participation rates actually did increase substantially for a long time. For example, Figure 3.2 shows that the female labor force participation rate at age 35 increased by about 30 percentage points between cohorts born in the early 1930s and those born in the early 1950s. A large literature explores many possible reasons for this trend, such as advances in home production technology (Goldin, 2006, for example) and diffusion of more effective contraception (Bailey, 2006; Goldin and Katz, 2002).

Closer inspection of Figures 3.1 and 3.2, however, poses a puzzle. Figure 3.2 shows that women's labor force participation rates, after rising for a long time, reached a plateau for birth cohorts from the early 1950s on. Although Figure 3.2 documents the plateau only for age 35, Fallick and Pingle (2009) and Lee (2012) show that the plateau applies more broadly. On the other hand, Figure 3.1 shows that women's college-going has continued to rise beyond the cohorts born in the early 1950s. So can changing expectations about future labor force attachment really explain the continued rise in women's college-going?

The answer is that it can if the trend across cohorts in expectations about future labor force attachment has lagged behind those cohorts' eventual realizations of labor force attachment in adulthood. For example, Figure 3.2 shows that women born in 1960 and women born in 1975 have turned out to have similar labor force participation rates. But what if the women born in 1960 did not anticipate when they were teenagers that they were going to work for pay so much, but the women born in 1975 did? Then the difference in their beliefs as teenagers about their future work still could have contributed to a difference in college-going.

Fortunately, this story about changing work expectations does not have to be left to speculation. Because various waves of the National Longitudinal Surveys have asked teenage women about their expectations regarding future work, it is possible to track trends in teenagers' beliefs across cohorts. The next section describes the NLS data in detail and demonstrates that young women's anticipations of future labor force attachment have indeed risen across cohorts. It also is possible within each NLS to test my model's prediction that those teenage women who express stronger expectations of future labor force attachment are more likely to go on to college. Section 3.4 will present the results of such tests.

3.3 Data

My data come from three National Longitudinal Surveys: the NLSYW68, which began by surveying 5,159 young women aged 14 to 24 in 1968; the NLSY79, which began by surveying 12,686 young men and women aged 14 to 22 in 1979; and the NLSY97, which began by surveying 8,984 young men and women aged 12 to 16 in 1997. Those surveys provide nationally representative data and collect detailed information on future work plans, educational attainment, and family background of the respondents.[2] Moreover, since each survey represents

[2] Unfortunately, data on future *work* plans of men comparable to those of women from the NLSYW68 are not available. The National Longitudinal Surveys of Young Men 1966 (NLSYM66) asked about future *job* plans under a presumption that all men would be working around age 30. To be specific, the question asked is "what kind of work would you like to be doing when you are 30 years old?" The answer should be either a specified job code or a choice between "same as present job" and "don't know."

different generations, data from the three surveys allow me to look at long-run trends in teenagers' work expectations for a broad range of birth cohorts.

To examine the trend in women's early work expectations across cohorts, I create a unified measure of their anticipated participation, a dummy variable indicating whether an individual expects to be working or not at a later age. I construct the work expectation measure based on information on their future work plans surveyed in early rounds of each survey. The NLSYW68 and the NLSY79 ask whether a respondent expects to be working or not at age 35. The original question is a multiple-choice question that reads: "what would you like to be doing when you are 35 years old?" The choices are 1) working; 2) same as present job; 3) don't know; 4) married, keeping house, raising family; 5) other. Although the code varies with each round, working and getting married are always included in the choice set. The respondents who said they planned to get married were asked about plans after marriage in a following question. The subsequent question is: "sometimes women decide to work after they have been married for a while. If you were to work, what kind of work would you prefer?" And the choices are 1) work (different job from present or last); 2) work (same job as present or last); 3) don't plan to work. Because those respondents chose marriage over work in the first question and the following question does not specify age at which they are working, I neglect the information from the subsequent question.[3]

[3] It seems conceivable that the first question does not provide useful information because the choices are not exclusive to each other. A respondent might expect to get married and to be working at age 35, but the question forces her to choose either of them. To examine the possibility, I compare women's responses to men's. While the fraction of women from the NLSY79 who expressed an expectation of working at age 35 varies across cohorts, that of men from the same sample was stable across cohorts with a rate higher than 0.9. Therefore, I believe

These data, however, are not available for the NLSY97. The NLSY97 asks instead the anticipated percentage chance that a respondent will be working 20 or more hours per week at age 30.[4] Figure 3.4 shows the amount of variation in the variable. The distribution is considerably right-skewed with a weighted average of 94. Of the respondents, about 81 percent at age 16 and 84 percent at age 17 answered that the chance would be greater than or equal to 90 percent. To construct an expectation variable similar to that from the NLSYW68 and the NLSY79, I recode the percentage rate and create a dummy variable that equals one if the rate is above 50 and zero otherwise. Due to the different context of the question and the arbitrariness of the variable construction, the work expectation data from the NLSY97 are not quite comparable to those from the NLSYW68 and the NLSY79. In any case, given the necessarily arbitrary wording of the survey questions, the measure of work anticipation from each of the three surveys is only a crude proxy for the young women's beliefs about their future labor force attachment.

Figure 3.3 displays the data by age and sample. The vertical axis represents the share of women who thought they would be working in their thirties. Note that point estimates for older women from the NLSYW68 and the NLSY79 are based on more observations due to the panel nature of the samples. For example, age 14 from the NLSYW68 contains women born in 1954 and no longer exists in the data after 1968. Age 19 from the same sample, however, includes women born in 1949 to 1954 because 19-year-old women are observed simultaneously in 1968 through 1973. In the case of the NLSY97, the ages 16 and 17 contain women born in 1980 and 1981, respectively, since the NLSY97 asked the question only at the first round of the

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the first question provides useful information unless women and men take the meaning of the question differently.

[4] The NLSY97 asks: "what is the percent chance that you will be working for pay more than 20 hours per week when you turn 30?"

interview.[5] The figure shows that there is a strong cohort effect, but not a noticeable age effect in the trend. In all age groups, the fraction of young women who expect to be working at age 35 rose sharply across cohorts – around 0.45 for the early 1950s cohorts and 0.75 for the early 1960s cohorts. The estimate for the early 1980s cohorts from the NLSY97, which is weakly comparable with those from the previous two cohort groups, is 0.95. This increasing pattern of women's anticipated labor force attachment and the upward trend in their college-going are consistent with the proposition that the two are positively related.

Figure 3.5 illustrates their positive relationship more clearly. It shows four trends: actual labor force participation in the broken black line; teenagers' anticipated labor force participation at age 35 in the red line with circle marks; teenagers' anticipated chance of working 20 or more hours of work at age 30 in the grey line with triangle marks; and college attendance rates in the solid blue line.[6] The rates of both college attendance and anticipated participation have risen across birth cohorts, while the actual participation rate has plateaued. The rise in women's college-going and the plateau in labor force participation happened around the same time because not all of the early work plans were realized. In fact, the discrepancy between the anticipated and actual participation is quite large. Women's actual labor force participation can be viewed as a realized version of young women's anticipated participation. The comparison of

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[5] The NLSYW68 and the NLSY79 surveyed future work plans repeatedly, from 1968 to 1987 for the NLSYW68 and from 1979 to 1984 for the NLSY79.

[6] An odd feature of Figure 3.5 is that the anticipated participation rate for the 1950 cohort is much lower than for any other cohort. It seems that this lower estimate is due to a mysterious 1968 survey effect. The anticipated participation rate from the 1968 wave is smaller than in any other wave not only for the 18-year-olds, but for other ages as well. The regression results in the next section are qualitatively robust to this survey effect. Excluding observations from the 1968 wave removes 370 observations from my regression sample, but the magnitude of the estimated effect of the anticipation variable on college-going increases by around 0.04, so that the significance level stays about the same.

the anticipated and actual participation reveals that while earlier cohorts of women considerably underestimated how much they would turn out to work, recent cohorts of women overestimated their future labor force participation.

The main thesis of this paper is that rising expectations of future labor force attachment among teenage women may play an important role in women's increasing college-going. This story involves two empirically testable propositions: 1) that teenage women's expectations of future work have indeed increased across cohorts, and 2) that teenage women who express stronger expectations of future work are more likely to attend and complete college. Having used the NLS data in this section to verify the first proposition, in the next section I will use them to test the second. Because the regression analyses in the next section will focus on the relationship between young women's anticipated labor force attachment and their college-going, I will restrict the regression samples to young women with information on their work expectations. For women from the NLSYW68 and the NLSY79, I use their work expectations when they are 18 since the majority of young women decide whether to go to college at that age.[7] For women from the NLSY97, I use their work expectations when they are 17 because their anticipation at age 18 is not available. I also limit the college outcome variables to enrollment by age 19 and completion by age 25 because the work anticipation when they are 17 or 18 is more likely to be relevant to college-going decisions made around age 18.

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[7] Among women with information on college enrollment by age 19, about 18 percent of women from the NLSYW68 and 21 percent of women from the NLSY79 are already in college at age 18. Empirical results in the next section are robust to using their work expectations when they are 17. Using the earlier work expectations measure, I lose at most 533 observations in my regression sample and the magnitude of the estimated effect of the expectations diminishes by about 60 percent although the effect still remains highly statistically significant.

I exclude observations with missing information on the college outcomes and family background from my regression samples. Some supplementary regressions will include controls for intelligence: IQ scores for the NLSYW68, percentile scores for the Armed Forces Qualifying Test (AFQT) for the NLSY79, and percentile scores on the Armed Services Vocational Aptitude Battery (ASVAB) for the NLSY97. However, the main regression samples will keep observations with the intelligence measures missing because the deletion of those observations leads to a sample that seriously over-represents college-goers. See Table 3.A1 for the details of how each deletion changes the mean of each college outcome variable.

In the end, the regression samples from the NLSYW68, the NLSY79, and the NLSY97 contain 1,734, 2,377, and 731 observations, respectively. Table 3.1 provides weighted summary statistics by sample. Women from the NLSY97 have the highest average college attendance and completion rates, while women from the NLSY79 have higher average rates than those from the NLSYW68, consistent with the national trends. However, as shown in Table 3.A1, all the estimates are a bit higher than the comparable estimates from the Current Population Surveys (CPS). Regarding college attendance, this discrepancy may come in part from the different definitions used in the two data sets. In the NLS samples, individuals who have enrolled in college are coded as college-goers, even individuals with less than one year of college experience. In the CPS, however, individuals with less than one year of college education are categorized as high school graduates. Because women who were 19 or older at the first interview do not have information on their anticipated labor force participation when they were 18, they are dropped from the regression samples. As a result, each sample consists of younger cohorts compared to the original sample. Whites are over-represented in the samples from NLSYW68 and NLSY79, but are under-represented in the sample from NLSY97.

3.4 Evidence on the association between work expectations and college-going

To examine whether teenage women who express higher expectations of future labor force attachment are more likely to attend and complete college, I estimate models of the form

$$\text{(6)} \qquad \Pr(D_i = 1) = F(a + bH_i + X_i c)$$

where D_i is a binary college-going outcome for woman i, $F'(\cdot) > 0$, H_i is a measure of the woman's anticipated labor force attachment, X_i is a vector of observed demographic and family background variables that may affect her college-going decision, and a, b, and c are parameters. The analysis in Section 3.2 predicts that b is positive. Due to data availability, H_i is measured in two ways. In regressions based on the NLSYW68 or NLSY79, H_i is a dummy variable that indicates whether the teenage woman expects to be working at age 35. In regressions based on the NLSY97, H_i is the teenage woman's anticipated chance of working 20 or more hours per week at age 30. In order to retain a sufficient sample size, I use a relatively sparse set of control variables. The vector X_i includes parents' years of schooling, dummy variables for white and birth year, and in some regressions, intelligence measures.

Tables 3.2, 3.3 and 3.4 present results from estimating the logit version of equation (6). The three tables are based respectively on samples from the NLSYW68, the NLSY79, and the NLSY97. The tables report the estimated average partial effect (APE) for each explanatory variable.[8] In columns (1)-(3), the dummy dependent variable is for attending college by age 19; in columns (4)-(6), it is for completing college by age 25. The specifications in columns (1) and (4) contain only H_i as an explanatory variable; those in columns (2) and (5) additionally include parents' years of schooling and dummies for white and birth cohort; those in columns (3) and (6) also control for an intelligence measure. In these tables, I report results based on the weighted quasi-maximum likelihood estimator (Manski and Lerman, 1977). Tables 3.A2-3.A4 show the logit coefficient estimates along with robust standard error estimates. The standard error estimates accompanying the estimated APEs in Tables 3.2-3.4 are calculated by the delta method (Bartus, 2005).

In Table 3.2, the estimated effect of the work expectations dummy from every specification is positive and statistically significant, which indicates that women who expect to be working at a later age are more likely to attend and complete college. The estimates in columns (2) and (5) mean that, other things equal, the college attendance and completion probabilities are about 0.17 and 0.14 higher, respectively, when a young woman expects to be working at age 35. Comparing across columns, once the demographic and family background variables and IQ scores are added as controls, the estimate becomes smaller in magnitude. Estimated effects of the control variables also show expected signs. As many studies point out, parents' schooling has a positive association with college attendance and completion.[9] Whites are less likely to invest in college education, other things equal, although the negative

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[8] The reported APE is the weighted sample mean of the estimated partial effects for each individual in the sample.

[9] For example, see Haveman and Wolfe (1995). However, there is an ongoing debate about the positive association. Behrman and Rosenzweig (2002) and Plug (2004) argue that an increase in the schooling of mothers would not have a positive impact on the schooling of children.

relationship is statistically significant only when controlling for IQ scores. The negative estimate on white is fairly common in the literature (for example, Jepson, 2008; Light and Strayer, 2002).

Table 3.3 presents results very similar to those in Table 3.2. Young women's expectations of working at a later age are highly correlated with their college-going outcomes in the NLSY79, as well as in the NLSYW68. One difference from Table 3.2 is that the estimated effects for the white dummy are always statistically significant, and even positive in the regressions for college completion. With AFQT percentile scores as controls, however, the estimate turns back to negative. Interestingly, the estimated effects for the work expectations variable are smaller in magnitude for college completion than for college attendance. One possible explanation is that the time gap between the moment at which expectations are reported and the time of college graduation may weaken the relationship. In my empirical setup, the gap could be seven years at maximum, which is a long enough period to change individuals' circumstances. Unforeseen events arising during the period, such as economic hardship or waning interest in school, may prevent some sample members from completing college.

Table 3.4 presents the results using the NLSY97 sample. The estimated effects of the work expectations variable are, again, positive and usually statistically significant. For interpreting the results in Table 3.4, on needs to recognize that the key explanatory variable is now the anticipated percent chance that an individual is working 20 or more hours at age 30. Unlike in Table 3.3, the estimated APEs are similar in magnitude across the college outcomes. The estimates in columns (2) and (5) mean that a 10 percentage point increase in the anticipated chance of work increases the college attendance and completion probabilities by about four percentage points. Once ASVAB percentile scores are added as controls, the estimates become less significant, maybe because of a small sample size caused by the deletion of observations with the percentile scores missing. Except for the estimates for father's years of schooling, estimates for other covariates are generally significant in a statistical sense. The magnitude and statistical significance of estimates on white vary widely across the specifications.

Further analyses reported in Tables 3.5 and 3.6 check the robustness of the results to weighting and specification of functional form. The weighted estimation may not be appropriate for my samples, especially the samples from the NLSYW68 and the NLSY79, in which blacks are under-represented, because the sampling weights provided by the National Longitudinal Surveys are to adjust for over-representation of blacks. Thus, in Table 3.5, I also perform unweighted estimation of the specifications from columns (2) and (5) of Tables 3.2-3.4. The unweighted estimates reported in columns (1) to (6) in the table indicate that there is no important difference between the weighted and unweighted estimates. The unweighted estimates for work expectation variables are generally a little smaller than the weighted estimates, except the one from the regression for college completion using the NLSYW68. The sign and significance of the unweighted estimates, however, are almost identical to those of the weighted estimates.

To check robustness to functional form, I estimate linear probability models of college-going, as an alternative to the logit specifications in the preceding tables. Columns (1) to (6) in Table 3.6 present the results. The weighted least squares (WLS) estimates in general agree in sign and significance with the logit APE estimates, with the WLS estimates being of a bit greater magnitude.

Combined with the evidence in Section 3.3 that young women’s expectations of future labor force attachment have increased across cohorts, this section’s evidence that such expectations are positively associated with college-going suggests that rising expectations of labor force attachment may account for at least some of the increase in women's college attendance and completion. But how much? To get a rough idea, I perform an Oaxaca-Blinder decomposition. I use the following simple decomposition technique:

$$(7)\quad \overline{Y^{97}} - \overline{Y^{68}} = \overline{X^{97}} b^{*97} - \overline{X^{68}} b^{*68}$$
$$= \underbrace{(\overline{X^{97}} - \overline{X^{68}}) b^{*68}}_{\text{explained effect}} + \underbrace{\overline{X^{97}} (b^{*97} - b^{*68})}_{\text{unexplained effect}} \quad ,$$

where $\overline{Y^{j}}$ is a college attendance or completion rate of sample *j*, $\overline{X^{j}}$ is an anticipated participation rate of sample *j*, and b^{*j} is a coefficient estimate from the linear regression of college outcome on anticipated participation using sample *j*. For simplicity, I am emphasizing the decomposition based on the linear probability model, but a more elaborate extension for the logit specification (Fairlie, 2005) yields very similar results. One problem is that the decomposition needs comparable explanatory variables across samples, which are not available in my data sets. To estimate b^{97}, I use a binary variable indicating whether, at age 17, a woman expects with at least 90 percent chance to be working 20 or more hours at age 30. Setting the threshold as high as 90 percent stacks the deck somewhat against finding a large role for changing expectations. Nevertheless, it turns out that the change in women's early work expectations between the NLSYW68 and the NLSY97 can explain 66 percent of the change in their college attendance and 69 percent of the change in their college completion.[10] Of course,

[10] The estimates of the explained effect are 0.074 (with standard error of 0.011) and 0.060 (with standard error of 0.009), respectively, for the changes in college attendance (0.112) and completion (0.087).

the Oaxaca-Blinder calculation may provide only a crude approximation of the connection between the trend in women's early work expectations and the trend in their college-going because it is based on crude measures of work expectations, which are not quite comparable across samples.

3.5 Summary

In recent decades, U.S. women's rates of college attendance and completion have overtaken men's, and are now considerably higher. This study follows up on a suggestion by Goldin, Katz, and Kuziemko (2006) that rising expectations of future labor force attachment among young women may account for part of the increase in women's college-going.

In Section 3.2, I formalized the insight from Goldin et al. by presenting a simple theoretical model illustrating the role of work expectations in optimal college-going decisions. In Section 3.3, I used information on teenage women's work expectations from a succession of National Longitudinal Surveys to verify that work expectations have indeed risen across cohorts. In Section 3.4, I used the NLS data to corroborate that young women with higher work expectations are considerably more likely to attend and complete college. In combination, the results indicate that young women's rising work expectations may indeed be an important part of the story behind their rising rates of college attendance and completion.

Table 3.1. Summary statistics

Women from	NLSYW68		NLSY79		NLSY97	
	Mean	(Std. dev.)	Mean	(Std. dev.)	Mean	(Std.dev.)
Dependent variable						
1 if						
attending college by age 19	0.475	(0.500)	0.497	(0.500)	0.587	(0.493)
completing college by age 25	0.215	(0.411)	0.237	(0.426)	0.302	(0.459)
Independent and control variables						
Father's years of schooling	10.9	(3.78)	11.7	(3.42)	13.9	(3.33)
Mother's years of schooling	11.1	(2.78)	11.6	(2.63)	13.2	(2.99)
Anticipated chance of work when 17					0.933	(0.161)
1 if						
expecting to work at age 35 when 18	0.458	(0.498)	0.698	(0.459)		
expecting to work at age 30 when 17					0.934	(0.249)
white	0.899	(0.301)	0.826	(0.379)	0.721	(0.462)
birth cohort						
1949	0.196	(0.397)				
1950	0.184	(0.388)				
1951	0.207	(0.405)				
1952	0.231	(0.422)				
1953	0.181	(0.385)				
1960			0.173	(0.378)		
1961			0.222	(0.416)		
1962			0.199	(0.399)		
1963			0.202	(0.402)		
1964			0.204	(0.403)		
1980					1.00	(0.00)

Table 3.1. (cont'd)

Women from	NLSYW68	NLSY79	NLSY97
Number of observations	1,734	2,377	731

Note: Custom weights offered by NLS Investigator are used. Parents' years of schooling greater than 18 in the NLSY79 and the NLSY97 samples are recoded as 18, in the same fashion as the NLSYW68. The work expectations variable from the NLSYW68 and the NLSY79 indicates whether a respondent expects to be working at age 35, while that from the NLSY97 indicates whether she expects to be working 20 or more hours per week at age 30 with a chance higher than 50 percent.

Table 3.2. Average partial effects from weighted logit regressions using NLSYW68

Women from NLSYW68	Dependent variable					
	Attending college by age 19			Completing college by age 25		
	(1)	(2)	(3)	(4)	(5)	(6)
Dummy variable for						
expecting to work at age 35 when 18	0.181**	0.172**	0.158**	0.147**	0.140**	0.133**
	(0.026)	(0.022)	(0.027)	(0.022)	(0.019)	(0.025)
Father's years of schooling		0.033**	0.039**		0.024**	0.019**
		(0.004)	(0.005)		(0.003)	(0.005)
Mother's years of schooling		0.055**	0.038**		0.022**	0.016**
		(0.005)	(0.007)		(0.005)	(0.006)
IQ score			0.010**			0.009**
			(0.001)			(0.001)
Dummy variable for						
white		-0.038	-0.159**		-0.011	-0.149**
		(0.034)	(0.039)		(0.027)	(0.049)
birth cohort	No	Yes	Yes	No	Yes	Yes
Number of observations	1,734	1,734	1,092	1,734	1,734	1,092

Note: Custom weights offered by NLS Investigator are used. Each point estimate is the average of the partial effect for units across the sample distribution of each covariate. Robust standard errors are in parentheses. * denotes p < 0.05 and ** represents p < 0.01.

Table 3.3. Average partial effects from weighted logit regressions using NLSY79

Women from NLSY79	Dependent variable					
	Attending college by age 19			Completing college by age 25		
	(1)	(2)	(3)	(4)	(5)	(6)
Dummy variable for						
expecting to work at age 35 when 18	0.203**	0.156**	0.121**	0.117**	0.074**	0.054**
	(0.027)	(0.023)	(0.022)	(0.022)	(0.017)	(0.015)
Father's years of schooling		0.036**	0.019**		0.026**	0.012**
		(0.004)	(0.004)		(0.003)	(0.003)
Mother's years of schooling		0.045**	0.033**		0.030**	0.020**
		(0.006)	(0.005)		(0.005)	(0.004)
Percentile score on AFQT			0.007**			0.005**
			(0.000)			(0.000)
Dummy variable for						
white		-0.075**	-0.213**		0.039*	-0.085**
		(0.022)	(0.020)		(0.017)	(0.020)
birth cohort	No	Yes	Yes	No	Yes	Yes
Number of observations	2,377	2,377	2,316	2,377	2,377	2,316

Note: Custom weights offered by NLS Investigator are used. Each point estimate is the average of the partial effect for units across the sample distribution of each covariate. Robust standard errors are in parentheses. * denotes $p < 0.05$ and ** represents $p < 0.01$.

Table 3.4. Average partial effects from weighted logit regressions using NLSY97

Women from NLSY97	Dependent variable					
	Attending college by age 19			Completing college by age 25		
	(1)	(2)	(3)	(4)	(5)	(6)
Anticipated chance of work at age 30 when 17	0.415**	0.372**	0.233	0.459**	0.370**	0.285*
	(0.124)	(0.125)	(0.160)	(0.148)	(0.129)	(0.128)
Father's years of schooling		0.009	0.004		0.012*	0.007
		(0.006)	(0.006)		(0.005)	(0.006)
Mother's years of schooling		0.033**	0.018**		0.028**	0.016*
		(0.007)	(0.007)		(0.007)	(0.008)
Percentile score on ASVAB			0.008**			0.006**
			(0.001)			(0.000)
Dummy variable for white		0.161**	-0.035		0.191**	0.046
		(0.039)	(0.039)		(0.031)	(0.037)
Number of observations	731	731	593	731	731	593

Note: Custom weights offered by NLS Investigator are used. Each point estimate is the average of the partial effect for units across the sample distribution of each covariate. Robust standard errors are in parentheses. * denotes p < 0.05 and ** represents p < 0.01.

Table 3.5. Estimates from unweighted logit estimation

	APE from unweighted logit estimation					
Dependent variable	Attending college by 19			Completing college by 25		
NLS sample	68	79	97	68	79	97
	(1)	(2)	(3)	(4)	(5)	(6)
Dummy variable for						
expecting to work at age 35 when 18	0.164** (0.022)	0.146** (0.021)		0.143** (0.019)	0.069** (0.015)	
Anticipated chance of work at age 30 when 17			0.372** (0.001)			0.316* (0.126)
Father's years of schooling	0.032** (0.003)	0.029** (0.003)	0.005 (0.005)	0.021** (0.003)	0.024** (0.003)	0.008 (0.005)
Mother's years of schooling	0.049** (0.005)	0.035** (0.004)	0.028** (0.006)	0.022** (0.004)	0.025** (0.004)	0.025** (0.006)
Dummy variable for						
white	-0.020 (0.027)	-0.100** (0.020)	0.171** (0.037)	-0.030 (0.025)	0.036* (0.016)	0.195** (0.031)
birth cohort	Yes	Yes	No	Yes	Yes	No
Constant						
Number of observations	1,734	2,377	731	1,734	2,377	731

Note: Robust standard errors are in parentheses. * denotes $p < 0.05$ and ** represents $p < 0.01$.

Table 3.6. Estimates from weighted least squares estimation of LPM

	Weighted least squares estimation of LPM					
Dependent variable	Attending college by 19			Completing college by 25		
NLS sample	68	79	97	68	79	97
	(1)	(2)	(3)	(4)	(5)	(6)
Dummy variable for						
expecting to work at age 35 when 18	0.181** (0.024)	0.168** (0.025)		0.150** (0.021)	0.094** (0.021)	
Anticipated chance of work at age 30 when 17			0.369** (0.001)			0.298** (0.086)
Father's years of schooling	0.032** (0.004)	0.037** (0.004)	0.009 (0.006)	0.023** (0.003)	0.031** (0.004)	0.012* (0.005)
Mother's years of schooling	0.050** (0.005)	0.040** (0.005)	0.033** (0.007)	0.021** (0.004)	0.028** (0.005)	0.030** (0.007)
Dummy variable for						
white	-0.029 (0.033)	-.086** (0.023)	0.164** (0.039)	-0.008 (0.024)	0.013 (0.018)	0.182** (0.033)
birth cohort	Yes	Yes	No	Yes	Yes	No
Constant	-0.452** (0.047)	- 0.426** (0.047)	-0.438** (0.146)	- 0.304** (0.043)	-0.517** (0.047)	-0.672** (0.126)
Number of observations	1,734	2,377	731	1,734	2,377	731

Note: Weighted least squares estimation of LPM uses custom weights offered by NLS investigator. Robust standard errors are in parentheses. * denotes $p < 0.05$ and ** represents $p < 0.01$.

Table 3.A1. Weighted averages of college outcome variables by sample

Sample of women	College attendance by 19 (CPS estimate)	College completion by 25 (CPS estimate)	Observation count [attd/cmplt]
1. Before any deletion			
NLSYW68	0.315 (0.355)	0.167 (0.175)	4,103/5,123
NLSY79	0.401 (0.402)	0.205 (0.220)	5,114/5,649
NLSY97	0.603 (0.580)	0.362 (0.329)	4,181/3,843
2. After deletion of observations with missing values of work expectation, college outcome, and parents' education (Samples presented in Table 1)			
NLSYW68	0.475 (0.373)	0.215 (0.204)	1,734/1,734
NLSY79	0.497 (0.410)	0.237 (0.226)	2,377/2,377
NLSY97	0.587 (0.580)	0.302 (0.329)	731/731
3. After the deletion of observations with missing values of work expectation, college outcome, parents' education, and intelligence measure			
NLSYW68	0.527 (0.373)	0.240 (0.204)	1,092/1,092
NLSY79	0.502 (0.410)	0.241 (0.226)	2,316/2,316
NLSY97	0.603 (0.580)	0.320 (0.329)	593/593

Note: The comparable estimates for the NLSYW68 (the NLSY79) in Panel 1 are based on women born between 1942 and 1953 (1957 and 1964), while those for the NLSYW68 (the NLSY79) in Panels 2 and 3 are based on women born between 1949 and 1953 (1960 and 1964). The CPS estimates of college attendance are based on women aged 20-24 with some college education and those of college completion are based on women aged 25-29 with a bachelor's degree. The comparable CPS estimates for the NLSY97 are identical throughout the panels because the estimates are based on a single birth cohort, the cohort 1980.

Table 3.A2. Coefficient estimates from weighted logit regressions using NLSYW68

Women from NLSYW68	Dependent variable					
	Attending college by age 19			Completing college by age 25		
	(1)	(2)	(3)	(4)	(5)	(6)
Dummy variable for						
expecting to work at age 35	0.732** (0.108)	0.962** (0.131)	0.988** (0.174)	0.879** (0.134)	1.025** (0.146)	0.974** (0.185)
Father's years of schooling		0.188** (0.025)	0.246** (0.034)		0.176** (0.027)	0.141** (0.035)
Mother's years of schooling		0.312** (0.035)	0.244** (0.044)		0.160** (0.035)	0.123** (0.046)
IQ score			0.065** (0.007)			0.066** (0.008)
Dummy variable for						
white		-0.218 (0.195)	-1.08** (0.294)		-0.077 (0.193)	-1.01** (0.313)
birth cohort	No	Yes	Yes	No	Yes	Yes
Constant	-0.438** (0.075)	-5.67** (0.411)	-11.5** (0.846)	-1.75** (0.102)	-5.48** (0.423)	-10.9** (0.855)
Number of observations	1,734	1,734	1,092	1,734	1,734	1,092

Note: Custom weights offered by NLS Investigator are used. Robust standard errors are in parentheses. * denotes $p < 0.05$ and ** represents $p < 0.01$.

Table 3.A3. Coefficient estimates from weighted logit regressions using NLSY79

Women from NLSY79	Dependent variable					
	Attending college by age 19			Completing college by age 25		
	(1)	(2)	(3)	(4)	(5)	(6)
Dummy variable for						
expecting to work at age 35	0.832** (0.114)	0.817** (0.121)	0.744** (0.138)	0.710** (0.150)	0.662** (0.157)	0.592** (0.170)
Father's years of schooling		0.188** (0.023)	0.119** (0.024)		0.220** (0.031)	0.129** (0.033)
Mother's years of schooling		0.240** (0.032)	0.203** (0.033)		0.250** (0.041)	0.211** (0.042)
Percentile score on AFQT			0.041** (0.003)			0.053** (0.004)
Dummy variable for						
white		-0.400** (0.120)	-1.41** (0.157)		0.336* (0.149)	-0.874** (0.208)
birth cohort	No	Yes	Yes	No	Yes	Yes
Constant	-0.600** (0.098)	-5.17** (0.378)	-5.14** (0.392)	-1.69** (0.133)	-7.77** (0.536)	-8.30** (0.577)
Number of observations	2,377	2,377	2,316	2,377	2,377	2,316

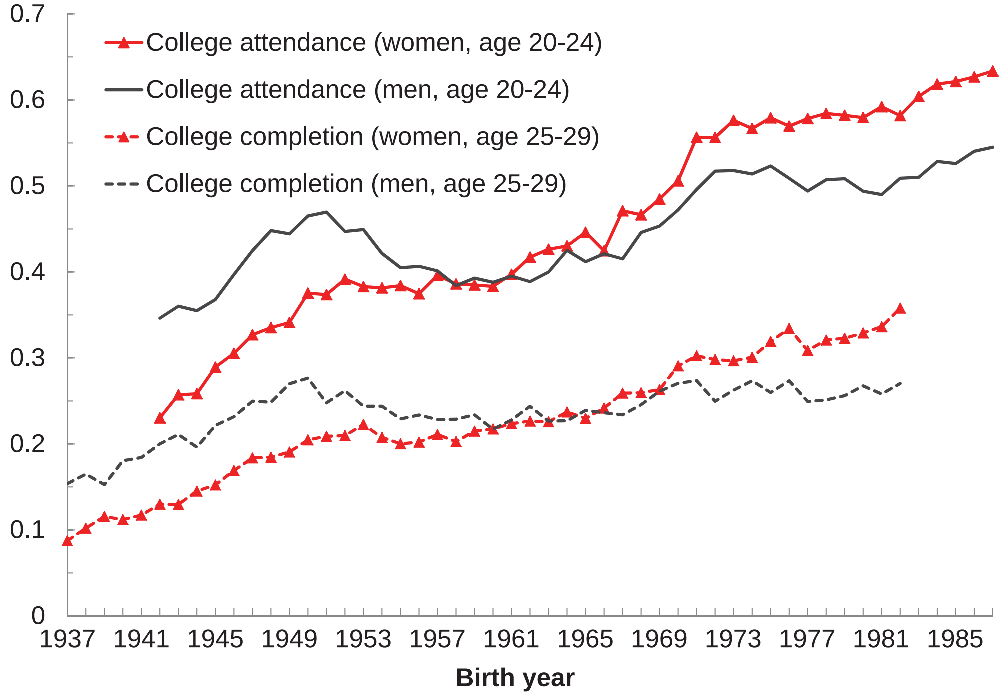
Note: Custom weights offered by NLS Investigator are used. Robust standard errors are in parentheses. * denotes $p < 0.05$ and ** represents $p < 0.01$.

Table 3.A4. Coefficient estimates from weighted logit regressions using NLSY97

Women from NLSY97	Dependent variable					
	Attending college by age 19			Completing college by age 25		
	(1)	(2)	(3)	(4)	(5)	(6)
Anticipated chance of work at age 30 when 17	1.74**	1.71**	1.44	2.23**	2.37**	2.20*
	(0.540)	(0.591)	(1.00)	(0.738)	(0.837)	(1.00)
Father's years of schooling		0.043	0.027		0.074*	0.056
		(0.029)	(0.036)		(0.033)	(0.043)
Mother's years of schooling		0.153**	0.113**		0.180**	0.124*
		(0.037)	(0.041)		(0.046)	(0.063)
Percentile score on ASVAB			0.051**			0.045
			(0.005)			(0.005)
Dummy variable for white		0.713**	-0.221		1.20**	0.353**
		(0.171)	(0.250)		(0.213)	(0.279)
Constant	-1.28	-4.35**	-5.01**	-2.96**	-7.52**	-8.21**
	(0.511)	(0.777)	(1.169)	(0.711)	(1.16)	(1.45)
Number of observations	731	731	593	731	731	593

Note: Custom weights offered by NLS Investigator are used. Each point estimate is the average of the partial effect for units across the sample distribution of each covariate. Robust standard errors are in parentheses. * denotes p < 0.05 and ** represents p < 0.01.

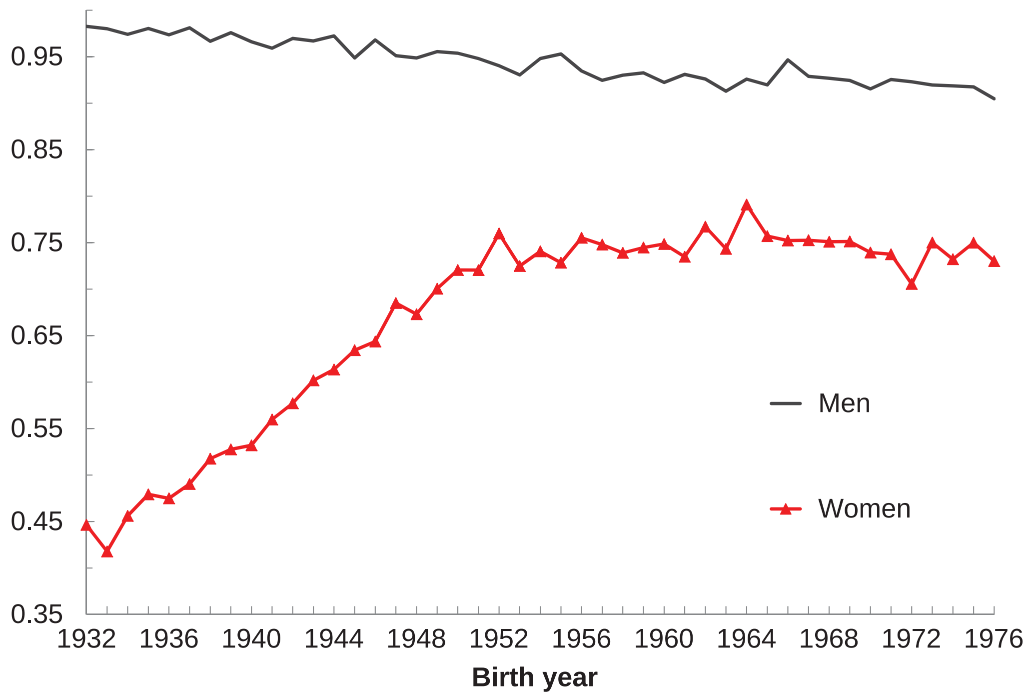
Figure 3.1: College attendance and completion rates



Note: CPS-provided sample weights are used. College attendance indicates at least one year of college education. College completion denotes four or more years of college education and/or the attainment of a bachelor's degree.

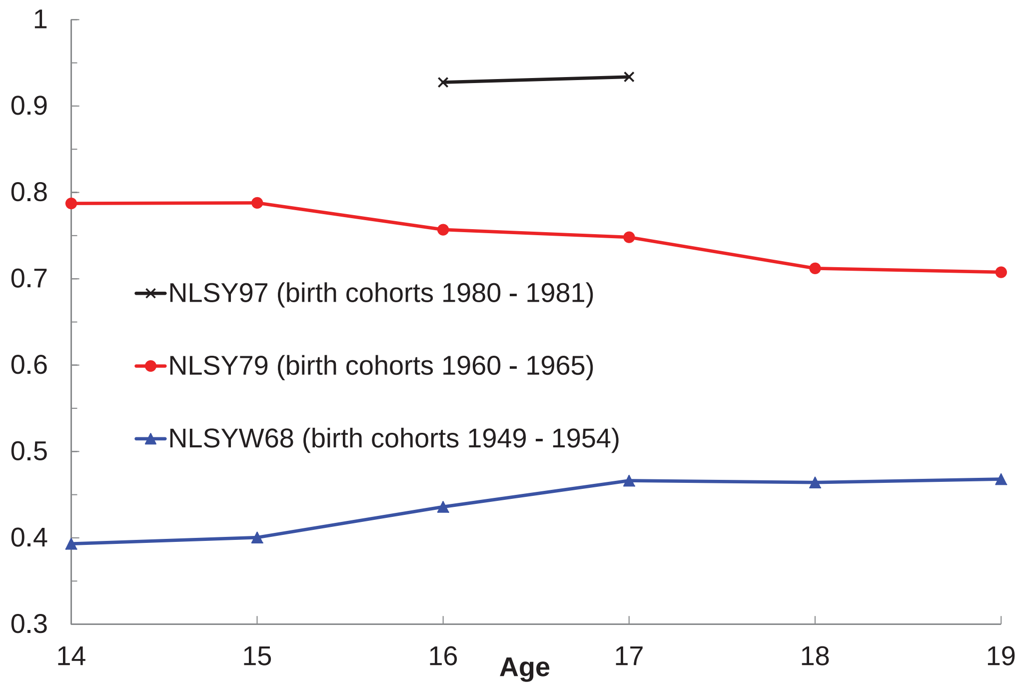
Source: 1962-2011 IPUMS-CPS.

Figure 3.2: Labor force participation rate, age 35



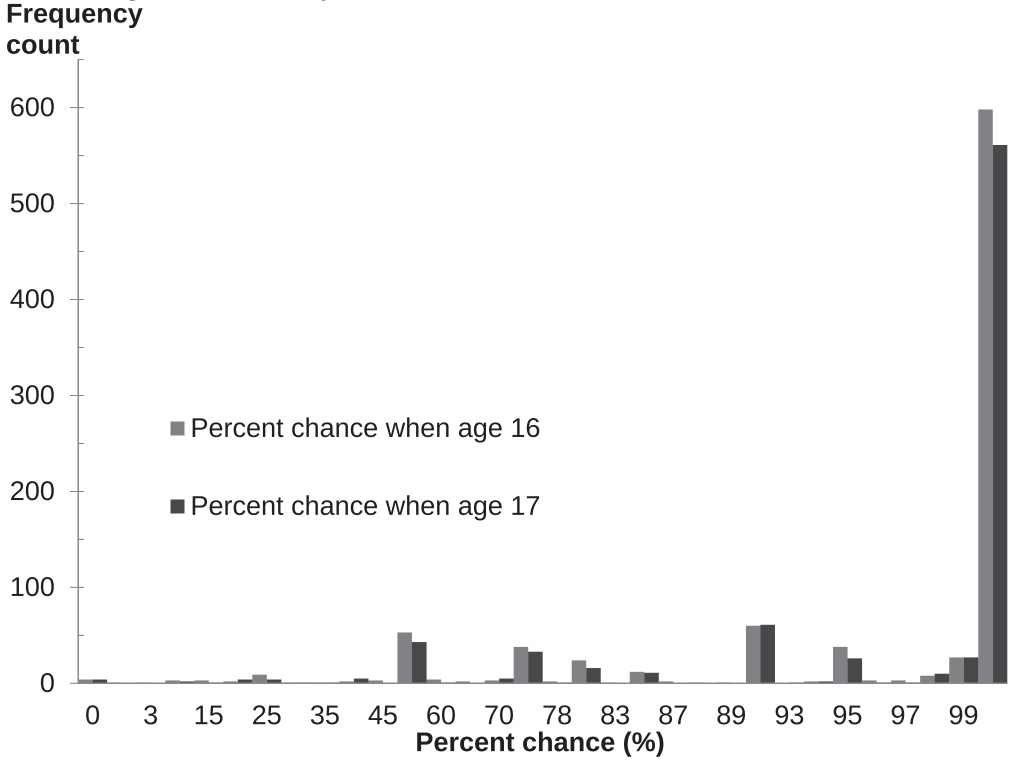
Note: CPS-provided weights are used.
Source: 1962-2011 IPUMS-CPS.

Figure 3.3: Women's work expectations



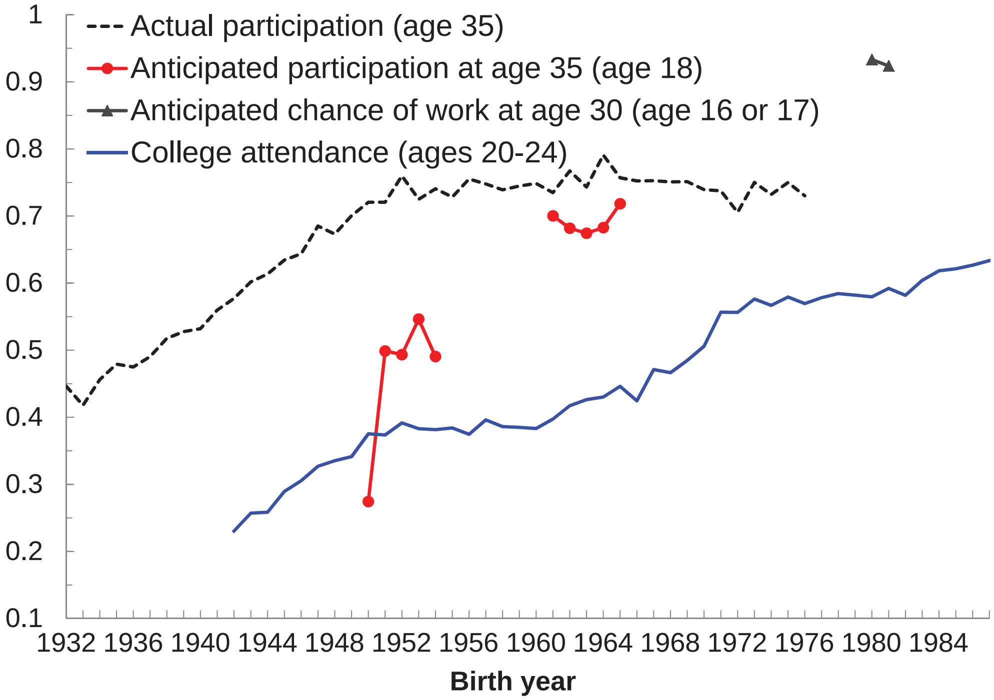
Note: The NLSYW68 and NLSY79 data are based on a question of whether a respondent expects to be working at age 35. The NLSY97 data are based on a question asked about the percent chance of working at least 20 hours per week at at age 30. I recode the chance and create a dummy variable which has 1 if the percentage rate is above 50 or 0 otherwise.
Source: The NLSYW68, NLSY79, and NLSY97.

Figure 3.4: Anticipated chance of work for women from NLSY97



Note: Total numbers of observations are 912 for age 16 and 822 for age 17.
Source: The NLSY97

Figure 3.5: Women's college attendance and anticipated and actual participation



Source: IPUMS-CPS (actual participation, college attendance), the NLSYW68 and the NLSY79 (anticipated participation), and the NLSY97 (anticipated chance of work)

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