Explaining the Returns to Delayed Childbearing for Working Women *

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Abstract

Using NLSY data, I observe a wage premium of 3% per year that women delay childbearing. For high-skilled women, the wage penalty from having children increases with time since the birth, but is more modest for delayers. I show strong correlations between age at first birth and observable characteristics such as education and experience. In a residual analysis framework, as much as 90% of the delay premium can be explained by differences in these characteristics for early and late child bearers. Some of the difference in experience is shown to be due to different leave-taking behavior around the first birth.

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1 Introduction

Since women began entering the work force in large numbers in the latter half of the twentieth century, they have faced important and difficult tradeoffs between family and work. One such tradeoff involves the timing of fertility. Biological constraints such as declining fecundity by age encourage early childbearing. However, researchers have found evidence of a wage premium to delay. Miller (2006) uses miscarriages as a source of variation in the timing of first birth, and finds an additional year of delay is associated with a 3% increase in wage rates and a 10% increase in earnings. She attributes the wage premium to differences in returns to experience for mothers, due to a "mommy track" career path. In a similar natural experiment design, Buckles (2007) shows that the ability to delay children using infertility treatments leads to delayed career interruptions, with some evidence of higher wages for mid-career women. Ellwood, Wilde, and Batchelder (2004) show that wage profiles for high-skilled women flatten out at the point of the first birth, resulting in significantly greater lifetime earning for delayers. Taniguchi (1999) uses NLSY data and finds that independent of work experience, delayed career interruptions are associated with higher wages—a result she attributes to missed "career building" opportunities for early child bearers. In Blackburn, Bloom, and Neumark (1993), women who have children later have higher early wages than women who have children earlier, which the authors attribute to greater human capital investment on the part of delayers.

Understanding the causes of this wage premium is essential to understanding the choices working women face and the decisions that they make. Is this premium due to discrimination against mothers, and early child bearers in particular? Are there features of the wage structure that reward delay? Perhaps women signal some aspect of their quality, such as ambition, by delaying childbirth? Or, is the premium generated by differences in human capital among early and late child bearers?

This paper will examine how much of the delay premium can be explained by the last of these hypotheses. First, I will measure the wage premium to delay, taking into consideration the different characteristics of women with first births at different ages. Using data from the National Longitudinal Survey of Youth, I find a raw return of approximately 3% per year of delay for hourly wages in 2003. Using the full NLSY panel and a fixed-effects framework, I show that there is a wage penalty to motherhood that increases with time since the birth, and that these penalties are greatest for high-skilled women. However, this effect is attenuated for women who delay childbirth.

Next, I show that there are significant differences in women's observable characteristics by age at first birth. Women who delay are more skilled, more educated, more likely to be in professional or managerial careers, and have more experience—even conditional on completed fertility. They also marry at a later age, and have fewer children. I return to estimating the return to delay, adjusting wages for differences in observable characteristics. I find that as much as 90% of the return to delay can be explained by differences in observable characteristics—reducing the wage premium from a year of delay from 3% to 0.3%. Education, experience, and age at first marriage have the most explanatory power. These results support a human capital theory for the delay premium.

Finally, I attempt to explain some of the experience difference between early and late child bearers by analyzing leave-taking behavior around the first birth. Delayers are less likely to take any leave, and those that do take leave are more likely to return to the same job. The average length of the gap for all women is shorter for delayers. These differences may be due in part to differences in maternity-leave coverage or access to child care.

This paper is organized as follows. In Section 2 I discuss evidence of delay behavior and returns to delay, and consider possible explanations for the return. Section 3 summarizes the data, and in Section 4 I document the return to delay. Section 5 describes the variation in observable characteristics of women with different ages at first birth, and examines how much of the delay premium can be explained by observables. In Section 6 I present data on leave-taking behavior. Section 7 concludes.

2 What explains delay?

The fact that natural fecundity declines by age makes delayed childbearing risky. Menken, Trussell, and Larsen (1986) find that, among women not using contraception, the percent of couples who remain childless varies by age at marriage: 6% for couples married at age 20-24, 9% at ages 25-29, 15% at ages 30-34, 30% at ages 35-39, and 64% at ages 40-44. As the probabilities of becoming pregnant and carrying a pregnancy to term fall, so does the probability of having a healthy child. Van Noord-Zaadstra, et al (1991) claim that the probability that a non-contracepting woman conceives and delivers a healthy baby falls by about half from age 25 to age 35.

Despite these risks, the increase in mean age at first birth in the United States has been well documented, by demographers and economists alike (see for example Mathews and Hamilton, 2002; Ventura et al, 2001; Heck et al, 1997; Goldin 1997; Caucutt, Guner, and Knowles, 2002). There are also clear relationships between delay and characteristics such as education and career attainment. Figure 1 shows the rise in mean age at first birth from 1982-1999, by education level. The increase in age at first birth is steepest for more educated women. Over this period the average age at first birth for all women increased by about two years, as within-group means increased and as more women attained higher levels of education. Figure 2 shows the distribution of the age at first birth for women in the NLSY in 2004, who have completed most of their childbearing. While the distribution is skewed to the right for all women, the distribution for high-skilled women is more symmetric with a higher mean. Below, I further analyze the relationship between women's characteristics and the timing of their first birth.

What explains the decision of many women to delay childbearing despite the risk? As discussed in the introduction, research has shown a wage premium to delayed childbearing that might motivate increases in the average age at first birth. Each of these papers attributes at least some of the delay premium to differences in human capital or in returns to human capital. Theoretically, Caucutt, Guner and Knowles (2002) generate delay behavior in a model with human capital accumulation and a marriage market when returns to experience are positive. Erosa, Fuster, and Rustuccia (2002) relate age at birth to job tenure, where early childbearing is costly because it increases the probability of future job separations. Other theoretical models that could generate a delay premium are a signaling model in which women signal their productivity to employers by delaying children, or a "threshold" model in which the career structure rewards uninterrupted or intensive human capital investments up to some pre-designated point in the career path.[1] Finally, discrimination against mothers (statistical or preference-based) could result in a delay premium.

If differences in human capital can explain the delay premium, the premium should be significantly reduced when controlling for differences in observable measures of human capital. Alternatively, if discriminatory behavior or signaling is driving the premium, a return to delay would still be observable after adding these controls. I now turn to the data to determine the role of observable characteristics in explaining the delay premium.

3 Data

The data used in this analysis is from the National Longitudinal Survey of Youth, 1979. The NLSY79 is a panel data set, with annual surveys done from 1979 to 1994, and biannual surveys from 1996 to 2004. The respondents were age 14 to 22 in 1979, and 39 to 48 in 2004, which is the last year available. Therefore, the data follows most women through the end of their fertile period. The full panel is used when constructing age-wage and experience-wage profiles, and when performing fixed- or random-effects regressions. It is also used to construct fertility histories and experience totals in all years. There were 6,283 women in the 1979 sample, and I restrict the sample to the 4,913 women who report having a birth at some point. Attrition reduces the sample to 3,365 by 2004. For results based on 2004 data alone, sample weights for that year are used.

Hourly wages are in year-2000 dollars, and are calculated as the total annual income from wages, salary, and tips, divided by hours worked in that year. Hourly wages of less than $1 per hour or more than $200 per hour are dropped from the sample, constituting at most 1.3% of the relevant sample. Actual experience in a given year is equal to the total number of hours worked from 1978 through that year, divided by 2000. AFQT percentiles are normalized by the year in which the individual took the AFQT test, which is used here as a measure of skill.

[1] Examples of thresholds in careers would be earning tenure for academics, finishing a residency for physicians, or making partner for lawyers.

4 Documenting the Return to Delay

I use NLSY panel data from 1979 to 2004 to document differences in women's earnings by age at first birth. Figure 3 shows that for women who worked in 2003, hourly wages are increasing in age at first birth. Empirically, the wage premium shown in this figure is estimated to be about 3% per year of delay. This result is not diminished by restricting the sample to women with two children, indicating that this effect can not be explained by higher fertility rates for early child bearers.

Figures 4 and 5 give a more complete picture of the effect of fertility timing on wages over the life cycle. These figures, along with Table 1, are an extension of work done by David Ellwood, Ty Wilde, and Lily Batchelder in their 2004 working paper, "The Mommy Track Divides: The Impact of Childbearing on Wages of Women of Differing Skill Levels." The data presented here allow the observation of women with later births, and for more post-birth wage observations for all women. Thus, I am able to present data for a wider range of timing behavior, including the important group of women who delay into their mid-thirties.

In Figure 4, high-skilled women are divided into groups by age at first birth, and age-wage profiles are plotted for each group. Several features of the data are worth mentioning. First, wages for these women are similar early in life for all groups. However, as age increases, women with early births move onto flatter profiles. In fact, the flattening of the profile seems to occur roughly around the age of birth. Furthermore, there does not seem to be any "catch-up" in wages, resulting in wider wage gaps late in life for early child bearers. Figure 5 presents the same data for low-skilled women, and while there are slight level differences in wages by age at first birth, we do not observe the flattening of the profile that the high-skilled women experience. It appears that the wage penalty to motherhood and the delay premium are greatest for high-skilled women.

Regressions confirming these results are summarized in Table 1. The results are coefficients from fixed effects regressions in which the sample is limited to working women with a first birth after age 21 who are 20 or older at the time of their interview, with hourly wages between $1 and $200. Women with births at 21 or earlier are dropped to capture women who are likely to have completed their education, so that their education decision can more plausibly be considered exogenous to the birth timing decision. The regressions are presented by skill group, and include controls for age. As no other controls are included at this point (beyond the fixed effects), Table 1 presents estimates of the return to delay, unconditional on endogenous factors such as experience, occupational choice, marital status, future fertility, and higher education. The important issue of controlling for these characteristics of the women will be addressed in the next section.

For each year of the panel, women are divided into categories based on the age of their first birth and the time since that birth. [2] The estimates should be interpreted as the effect of being in the given age-at-birth/time-since-birth category, relative to women of the same age who have not yet had a child.[3] Thus, the number "-10.40" in the upper left-hand corner of the table means that women who have a first birth before age 27 and who are in the first four years after that birth have wages that are 10.40% lower than women of the same age who have not yet had a birth.

There are several meaningful comparisons in the table. First, to compare women who are in the same age-at-birth/time-since-birth category but with different skill levels (looking across rows), we see that the wage penalties are greatest for high-skilled women. For example, high-skilled women with a first birth before age 32 have wage penalties that are at least two times larger than low-skilled women in the first 17 years after birth.

Second, comparing coefficients within a skill and age-at-birth category (looking down columns, within each group), we see the increasing penalties to motherhood or flattening of the wage profile suggested by Figure 4. This effect is most dramatic for high-skilled women, where for early child bearers the penalty increases from 18.69% in the first four years after the birth, to 32.58% in years 5 to 9, to 51.86% in years 10 to 17, and finally to 61.89% after year 18. A similar pattern is observed for women with later births.

Third, we can compare women with the same skill level and time since first birth, who had a first birth at different ages (comparing across groups). High-skilled women

[2] The figures in Table 1 are the percentage changes in wages. These coefficients used to create these figures are available in Appendix Table 1.

[3] In each year of the panel, the omitted case is women who have not yet had a birth. Because all women in the sample eventually have a child, the equation is not identified for the 2004 panel.

experience smaller wage penalties for a given time since birth if that birth occurred after age 31. For example, high-skilled women in years 0-4 after the birth have an 18.69% penalty if that birth occurred before age 27, but only a 10.10% penalty if it occurred after age 31. This return to delay for high-skilled women persists as years since birth increase, though the effect is smaller. For low-skilled women, on the other hand, we actually observe a delay penalty.[4] Both this effect and the delay premium for high-skilled women are likely due in part to differences in the endogenous variables mentioned above, such as actual experience, occupational choice, and household characteristics. These differences are addressed in the next section.

5 Age at First Birth and Observable Characteristics

Analysis

Figures 4 and 5 and Table 1 illustrate a flattening of the wage profile after the first birth and wage penalties for motherhood that are greatest for high-skilled women, though these effects are attenuated when women delay childbirth. But it is reasonable to think that women who choose to delay childbirth might make other choices as well. How do the observable characteristics of women vary by age at first birth? Figures 6 through 11 show differences in human capital, family status, and occupational choice for women with first births at different ages. The samples are taken from the 2004 wave of the NLSY79, and include women who had a first birth between the ages 16 and 36. Women with first births after this point are excluded due to small sample sizes.

Figure 6 indicates that women with first births at a later age are more educated and more skilled, as measured by AFQT scores. A one-year increase in age at first birth is associated with an increase of 2.07 in AFQT percentile and a 0.20 year increase in education. Figures 7 further illustrates the importance of education in fertility decisions. By age 24, over 80% of those with less than 12 years of education and over 50% of those with 12 to 15 years of education have already had a child, while only 12% of those with a

[4] This result is consistent with previous research which shows that low-skilled women with a teenage birth have higher wages than similar women who do not (Hotz, McElroy, and Sanders 1996; Hotz, Mullin, and Sanders 1997).

college degree have done so. There is some catch-up by the mid-thirties, though levels of childbearing remain lower for the more educated, as more educated women also have fewer children at any given age.

Figures 8 and 9 show differences in marriage behavior by age at first birth. In Figure 8, we see that women who delay children are more likely to be married in 2004, though some of this result can be attributed to the fact that these women married more recently. Figure 9 shows that women with a first birth at a later age also get married later, though the slope is less than one. We would expect a slope of one if women waited the same amount of time to have children after marrying, on average, regardless of the age at which they marry. This suggests that there is some variation in the timing of first birth that is not determined by age at first marriage. For women who have a first birth more than eight months after their first marriage, the time to first birth ranges from 5 years to less than two. The time to first birth is maximized by those who marry at age 21, and gradually decreases for later marriages. However, even women who delay marriage do not, on average, have children immediately. While the marriage decision is obviously an important component of the birth-timing decision, there is significant variation in age at first birth conditional on age at first marriage.

In Figure 10, we see that women who delay children have accumulated more experience by 2004. While some of this difference is because early child bearers subsequently have more children, there is still an increase in experience by age at first birth *conditional* on completed fertility. This fact is more surprising when one considers that delayers have less potential experience on average, as they are more likely to have gone to college. The plot of higher education plus years of experience illustrates this. Finally, Figure 11 shows differences in occupational choice by age at first birth. Primarily, delayers are more likely to be in professional and managerial careers.

These figures illustrate important differences in observable characteristics among women with first births at different ages. I now use residual analysis to determine how much of the returns to delay discussed above can be explained by differences in these characteristics. The results are the coefficients from a regression of the residuals from log wage equations on age at first birth. Intuitively, the residuals are log wages, risk-adjusted for observable characteristics. The coefficients in the table indicate how much of the risk-adjusted wage can be explained by age at first birth.

The results in Table 2 are based on residuals from random effects regressions using the NLSY 1979-2004. The sample is again limited to women with a first birth after age 21 who are 20 or older at the time of the interview and hourly wages between $1 and $200. I chose a random effects model because I wanted to add controls for constant variables such as race and AFQT score; Hausman tests fail to reject equivalence of fixed effects and random effects models.

Looking at the results for all women in Table 2a, we see a return of 3% per year of delay when wages are only adjusted for the woman's age. I interpret this as the raw return to delay, and this is the number cited in Section 4. I now add other controls to adjust the wages, adding variables in order of plausible exogeneity. Adjusting wages for race and AFQT score reduces the premium by about one-third; further adjusting for education decreases the remaining premium by one-half to 1.1%.[5] Adding family controls reduces the premium to 0.5%, and observable experience and occupation categories decrease it to 0.3%, or one-tenth of the raw return to delay.[6] This is a key result—adjusting wages for observable characteristics can explain 90% of the observed return to delay.

While this important result does not depend on the order in which the wage adjustment is done, the order does matter when thinking about the relative contribution of each variable in explaining the return to delay. Education decreases the remaining effect by 48% and therefore has the biggest percentage effect. Experience has a small effect on percentage points explained, but decreases the remaining gap by 40%. In Table 2b, I reverse the order in which the adjustment is done, and find that AFQT score and race have very little impact once other controls are added. Education, experience, and age at first marriage again have the biggest impact on any remaining gap. This is particularly true for high-skilled women, who experience the largest delay premium. It appears that differences in these characteristics by age at first birth explain most of the return to delay.

[5] Adriana Lleras-Muney (2005) finds that OLS and IV estimates of education do not yield significantly different results, suggesting that education is exogenous.

[6] In regressions in which occupation and experience are added separately, experience plays a larger role in explaining the wage premium than occupation. This is also the case when the order of the adjustment is reversed, as in Table 2b.

These results are presented visually in Figure 12, where adjusted wages are shown as a function of age at first birth. As more controls are used in the wage adjustments, the slope of the line becomes flatter. While the line with unadjusted wages is obviously positively sloped, the line with the full set of controls, Adj5, is almost flat.

The results in Table 3 are based on the same method as those in Table 2, but use only 2003 wages and OLS regressions to create the adjusted wages. These results capture the effect of age at first birth on wages at the end of the fertile period, and do not include random effects. Here, in the results for all women, 80% of the return to delay for women age 22 to 36 is explained by observable characteristics. The return to delay that remains after the full set of controls is added is not significantly different from zero, though the standard errors are larger than in Table 2 due to decreased sample size. Again, the measured return to delay is greater for high-skilled women. Both Table 3a and 3b indicate that education and experience have the most power in explaining the return to delay.

Discussion

The finding that the observed return to delay can largely be explained by observable characteristics, and by differences in education and experience in particular, supports a human capital-based model of the return to delay. If wage discrimination against mothers was significant, than a substantial motherhood penalty would remain after controlling for these characteristics. Women may find this result encouraging, as it suggests that a woman who is able to maintain high levels of labor supply after the birth of a child will not suffer a large wage penalty. However, discrimination against mothers could still exist, if employers offer mothers fewer hours, fewer promotion opportunities (which would affect observed occupation), or if mothers were less likely to be hired at all.

It is important to reconcile the above findings with the results of Miller (2006) and Buckles (2007). Using natural experiments that create exogenous variation in the age at first birth, both of these authors observe a real return to delay. Meanwhile, the above results suggest that we should not observe a practically significant real return as a result of exogenous variation in age at first birth, *if* all of the observed characteristics controlled for here are also exogenous to the birth timing decision. Of course, they are not—completed fertility, marital status, occupational choice, experience, and possibly education could all be affected by age at first birth. For example, a woman with an exogenous increase in her age at first birth may become more attached to her employer or to the labor force in that time, and thus her future childbirth-related gaps in employment may be shorter. She may also find that the exogenous delay results in lower completed fertility. Thus, exogenous shifts in fertility might generate a wage premium to delay if these other outcomes are also affected. This paper has explicitly controlled for these observable characteristics, and a much smaller wage premium is observed.

One limitation of the residual analysis approach used in this paper is that the results may be biased if women make career, family, or human capital investment decisions in anticipation of early or late childbearing. For example, a woman who plans to have children early in life might avoid occupations that penalize that behavior, or under-invest in education if she anticipates that early childbearing will reduce the returns to that education. In this case, estimates of the return to delay conditional on occupation and education would be biased downward. Miller's use of miscarriages as a natural experiment is helpful here, as miscarriages provide an unanticipated change in birth timing. The fact that Miller finds the same 3% wage premium using this technique that I find here *before* adding any covariates suggests that this bias is not practically significant. In the end, both the natural experiment approach and the residual analysis technique used here are valuable in understanding the return to delayed childbearing.

Finally, it is also important to consider the possible impact of selection into the labor force on the above findings. Among women in the sample who have had their first child within the last two years, 34.6% have a missing wage observation. Correcting for selection is a difficult task, made more difficult here because assumptions required by standard techniques may not apply.[7] However, one can consider the likely direction of the bias. If high-earning delayers are less likely to drop out of the labor force after a birth, the return to delay will be overstated. Alternatively, if these high-earning women

[7] One possible method would be median regression analysis, in which individuals with no reported wages are assumed to have wages below the median. However, this assumption likely does not hold here, as some high-earning women leave the labor force after the birth of a child. Similarly, a Heckman correction would require a variable that is exogenous to the wage equation, but not to the labor supply equation. The presence of young children is often used, but is clearly inappropriate here.

are more likely to drop out (because they have high-earning spouses, or because of greater accumulated wealth), the return to delay would be understated.

Claudia Goldin's finding (2006) that relatively few high-earning women drop out of the labor force for long periods of time supports the first hypothesis. Appendix Table 2 confirms this. The sample here is women who report having a first birth within the last two years. The results are the marginal effects from a probit regression, where the independent variable of interest is the most recent wage observation. The regressions also include a control for additional children born since the first birth, and for the number of years since the last wage observation. The results show that women with higher pre-birth wages are less likely to have a missing wage observation. If women are positively selected into the labor force on the basis of earnings potential, estimates of a family gap would be overstated. However, the coefficients for delayers are not statistically different from those for early and middle-aged child bearers. This suggests that the selection effect does not differ much by age at first birth, so estimates of the return to delay should be little affected by selection.

6 Differences in Leave-Taking Behavior

The previous section showed that the large differences in accumulated experience by age at first birth seen in Figure 10 are partly responsible for the observed wage premium for delay. These experience differences are not mechanical—if two women have the same number of children but at different ages, their accumulated experience after completing fertility would not be different if leave-taking behavior was the same and each worked the same number of hours when working. However, because mothers are more likely to work part-time, early child bearers may have fewer years of full-time work, leading to less experience at a given age. Leave-taking behavior may also differ, although the effect of delay is ambiguous. If delayers are more attached to the labor force or to their employers, delayers will be less likely to interrupt their careers and any interruptions will be shorter. Also, delayers may have better access to child care (either employer-provided or purchased), facilitating a quicker return to work. On the other hand, early child bearers may be more credit-constrained and unable to leave work for any substantial amount of time.

Figure 13 shows that delayers are much less likely to have an employment gap around the first birth. Of the leaves they do take, delayers are less likely to have a between-job gap, in which the woman leaves one job and returns to another employer after the birth, and more likely to return to the same employer. This feature might be important in future research in explaining the return to delay, since delayers who return to the same employer do not lose tenure and are less likely to take a job at lower pay upon return.

The average duration of first-birth related gaps is shown in Figure 14. When including only women who take leave, delayers actually have longer average gaps. However, when including women with no leave as zeroes, generating an “expected” gap by age, the average gap is smaller for delayers. Comparing this average for women with a first birth at 24 to those at 34, there is a decrease of about 30 weeks in the expected gap. The difference is about 20 weeks when limiting the sample to those that return in five years. If this behavior is repeated on subsequent births, the accumulated differences in experience could be quite large.

Finally, Figure 15 suggests that a small part of the differences in leave-taking behavior might be due to differences in employer-provided benefits. Delayers might have better jobs or longer or more permanent relationships with their employers, which could mean better benefits. Also, delayers had their births in later years, when provision of these benefits was more common. The 1994 Family and Medical Leave Act, for example, would have covered some delayers but not the early child bearers. Access to these benefits could explain why delayers were more likely to return to the same job after a birth-related gap.

7 Conclusions

This paper has documented a wage penalty to motherhood that increases with time since the first birth and is more severe for high-skilled women. I then show important differences in the observable characteristics of women by age at first birth, including increases in experience, education, skill, and age at first marriage for women who delay. I find a 3% raw return to delay that is almost completely explained by differences in observable characteristics, most notably education and experience. This result suggests that human capital theory is more important than discrimination in explaining the return to delay. Last, I show that some of the experience differential between early and late child bearers is due to differences in leave-taking behavior. This finding has important policy implications—more comprehensive maternity leave or child care provision might allow early child bearers to return to the work force, which would increase their experience level and decrease the return to delay.

Although the return to delay can be explained by observable characteristics, we should not conclude that the return is not important or that the question of why women delay is answered. For one thing, most of the observed characteristics are at least partially endogenous—number of children, age at first marriage, occupational choice, and experience are all affected by age at first birth. In the case of education, for example, it is not clear why delayers have higher education levels, or if there is any causal relationship. It may be that women who would like to delay are able to attain higher levels of education, or it could be that more education encourages delay. If it is the latter, what features of the labor market make delay beneficial for more educated women? Are the returns to education or experience different for mothers and non-mothers? Ongoing research suggests they may be.

Finally, while these results indicate that the average return to delay can be explained by observable characteristics, there may be groups of women for whom this is not the case. In particular, there may be certain types of career structures that elicit a return to delay for women with similar observables. Careers with a threshold promotion structure, such as academics, law, or medicine, might reward early intensive human capital investments more than later investments. These questions are left to ongoing and future research.

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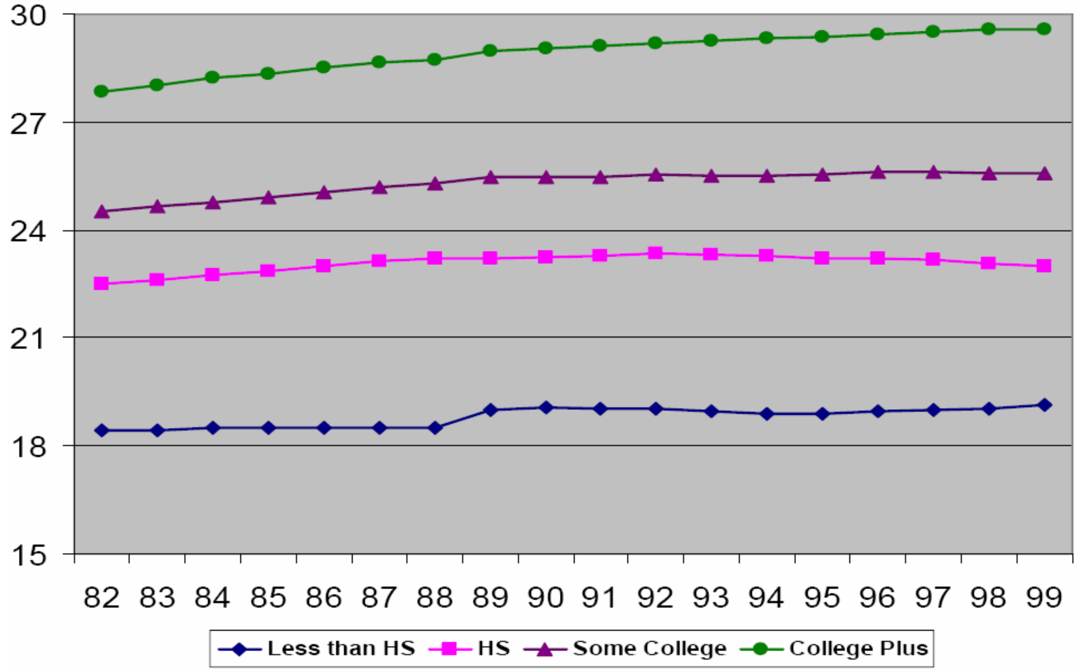
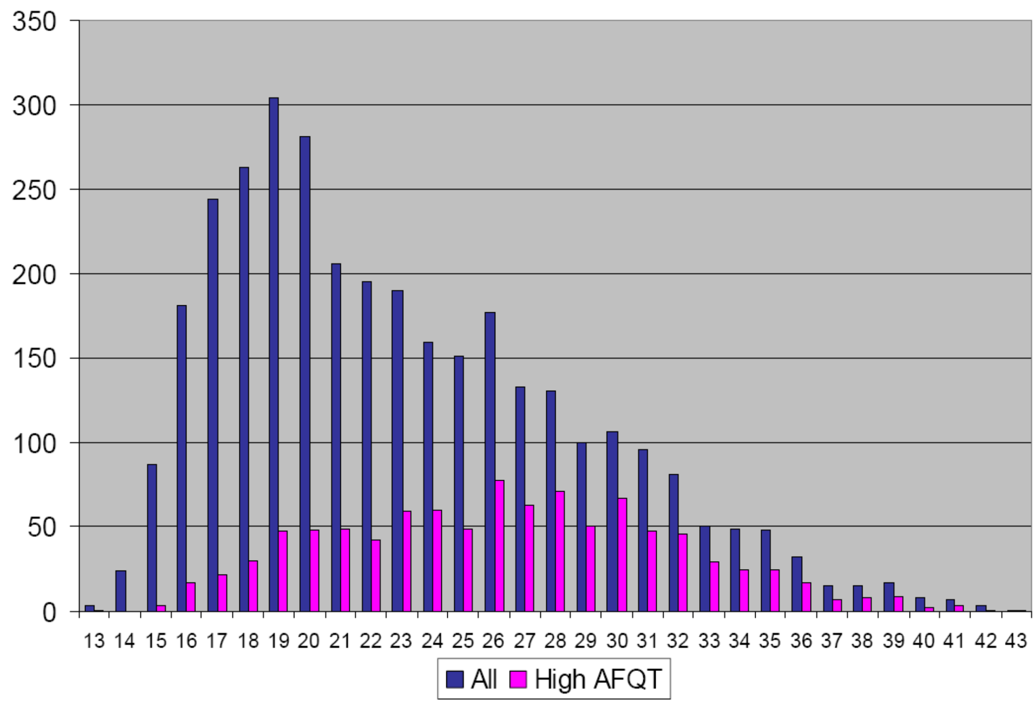
Figure 1: Mean Age at First Birth by Education, 1982-1999



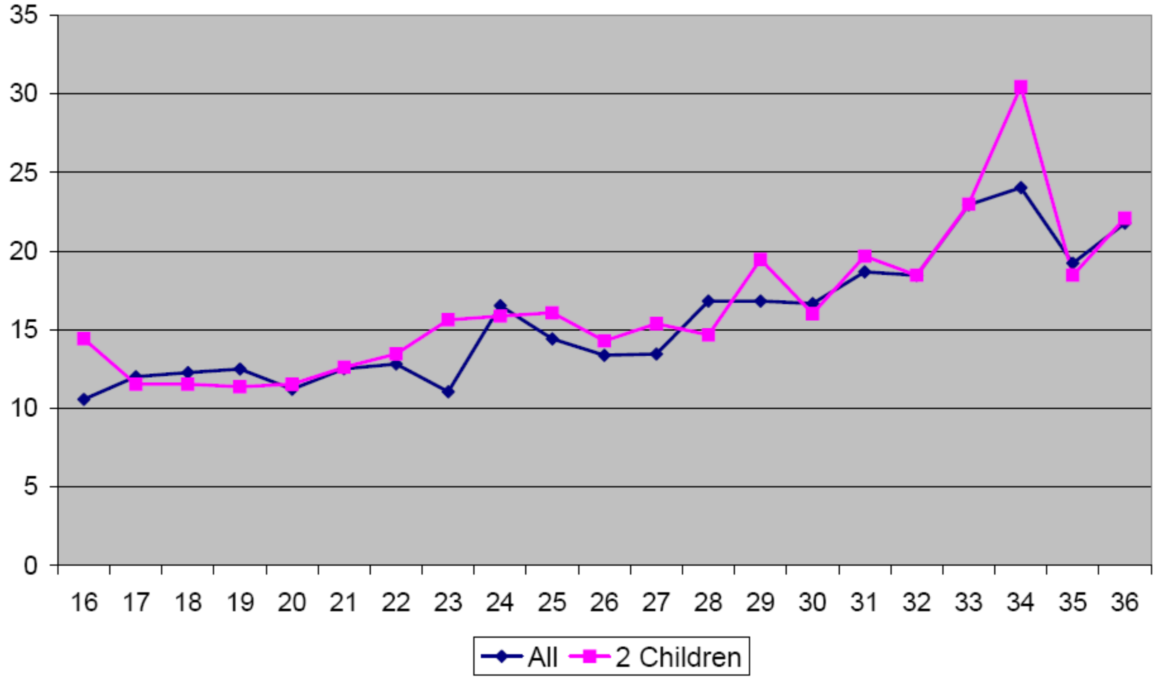
Source: Natality Detail Files, 1982-1999. Mothers' education is not required on all birth certificates until 1989. Births without this information are not included in the table.

Figure 2: Distribution of Age at First Birth, NLSY 2004



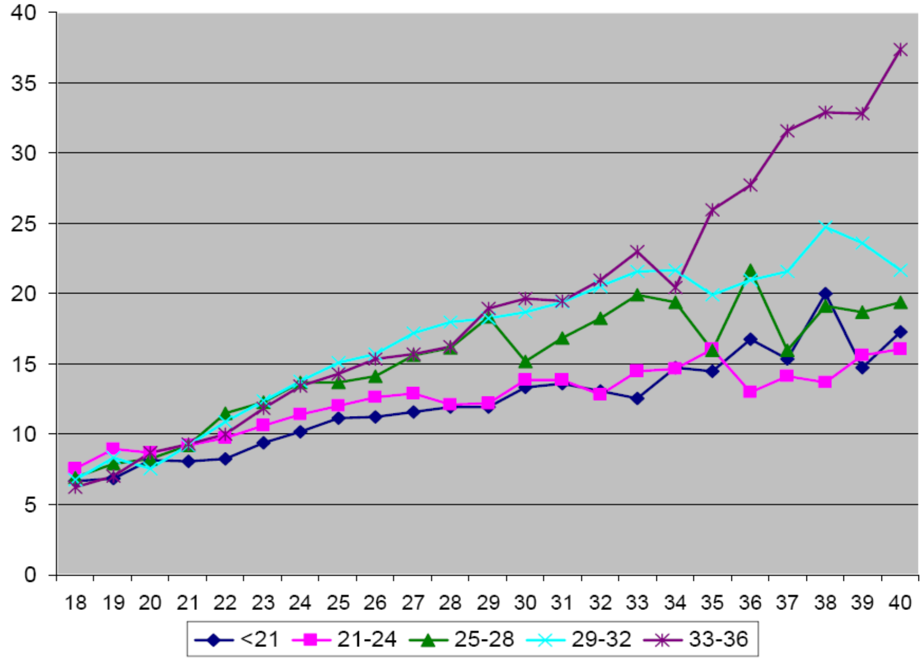
Source: National Longitudinal Survey of Youth, 2004. A women is classified as having a high AFQT score if her percentile score is in the top third of women who took the test at the same age.

Figure 3: Hourly Wages in 2003, by Age at First Birth



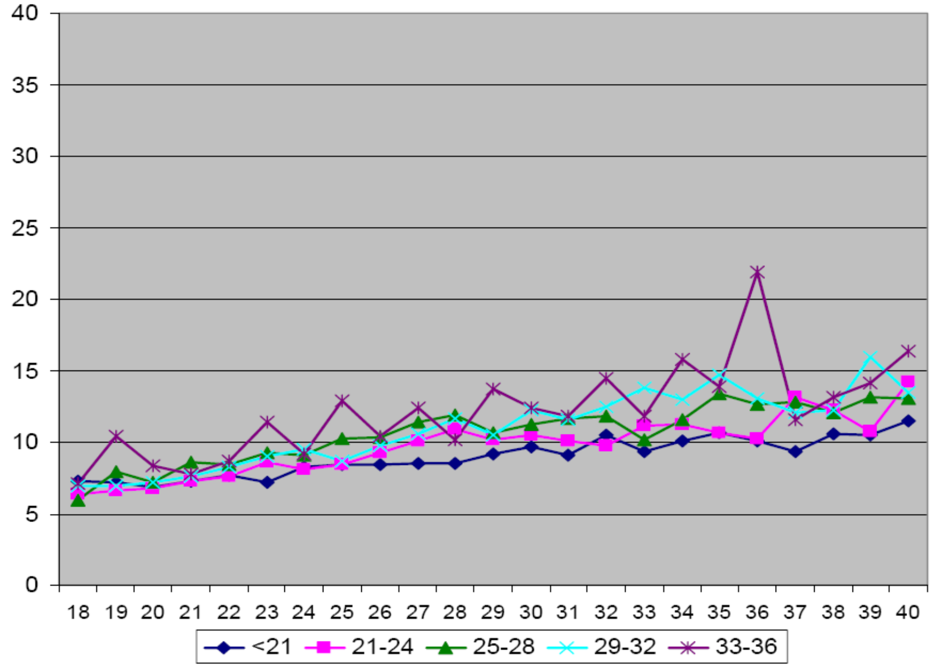
Source: National Longitudinal Survey of Youth, 2004. Hourly wages are calculated as total income from wages, salary, and tips in 2003, divided by total hours worked in 2003.

Figure 4: Age-Wage Profiles by Age at First Birth, High AFQT Women



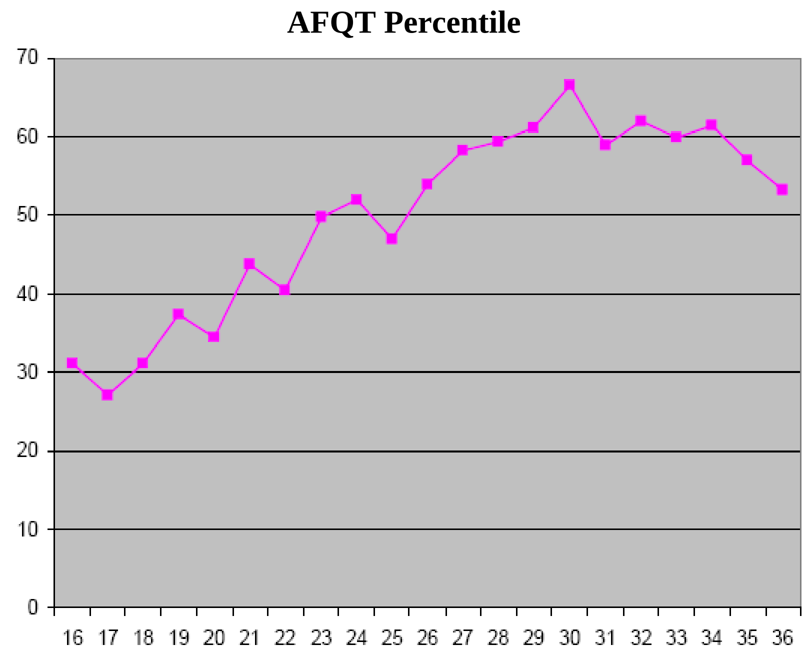
Source: National Longitudinal Survey of Youth, 1979-2004. A woman is classified as having a high AFQT score if her percentile score is in the top third of women who took the test at the same age. Hourly wages are in year-2000 dollars.

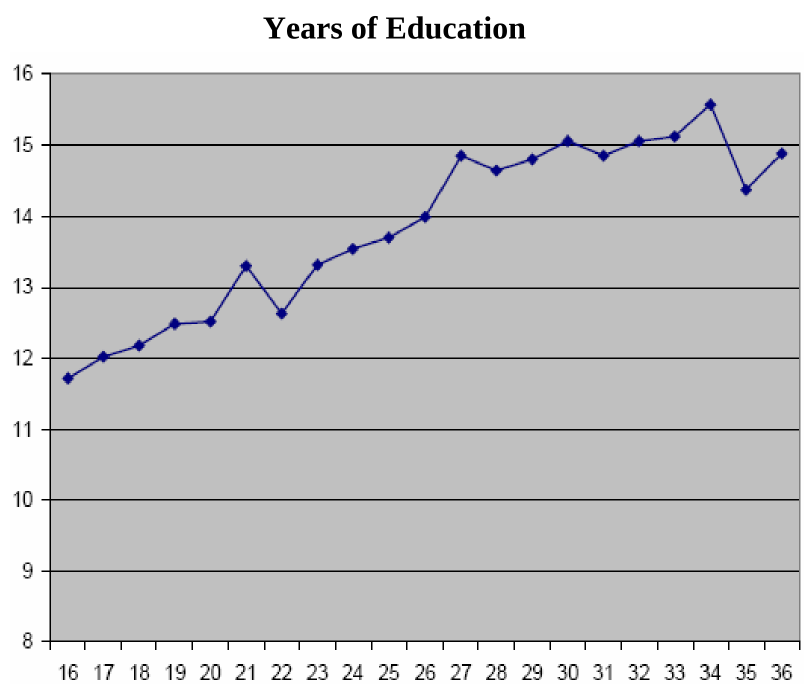
Figure 5: Age-Wage Profiles by Age at First Birth, Low AFQT Women



Source: National Longitudinal Survey of Youth, 1979-2004. A women is classified as having a low AFQT score if her percentile score is in the bottom third of women who took the test at the same age. Hourly wages are in year-2000 dollars.

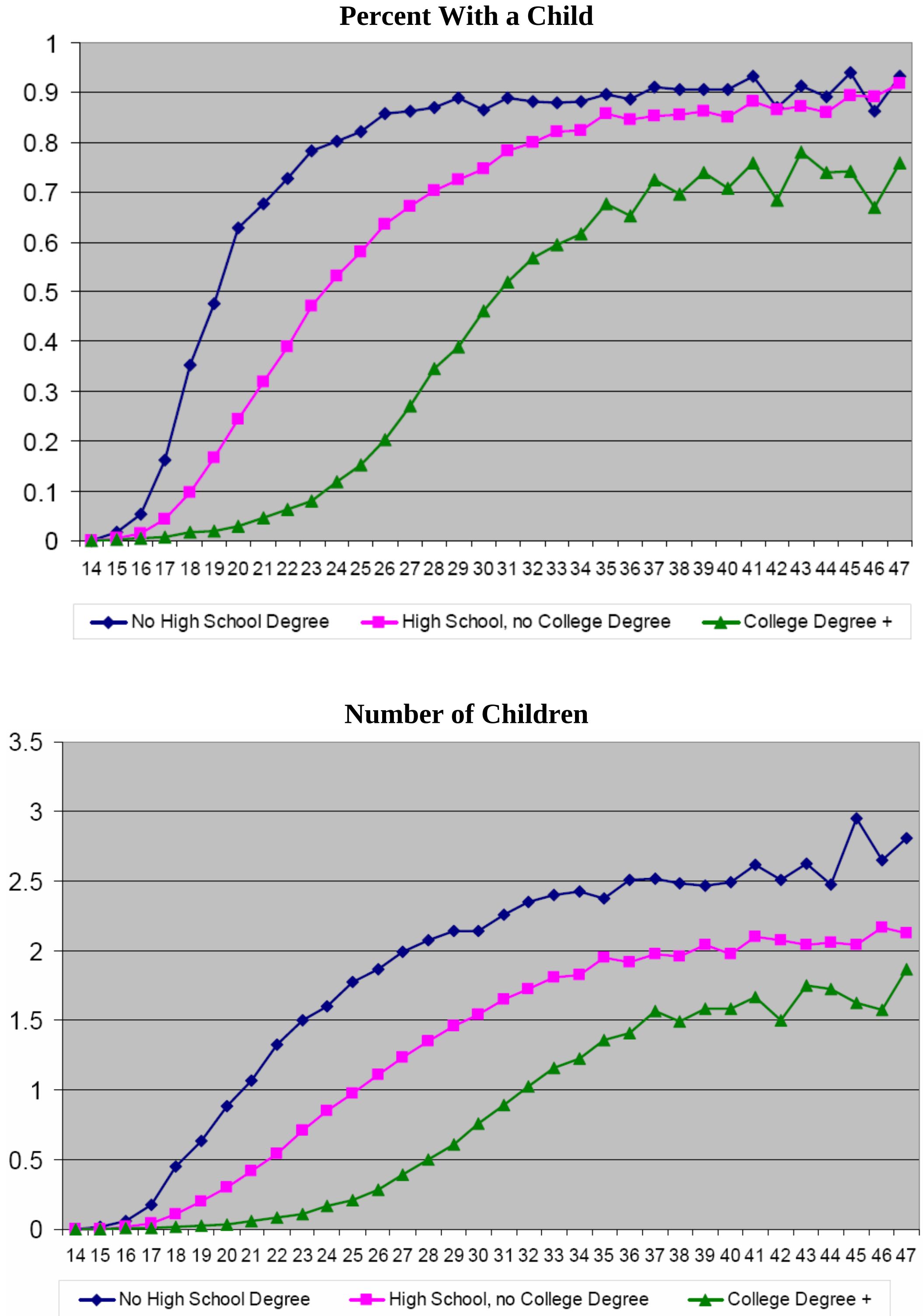
Figure 6: Mean Characteristics by Age at First Birth





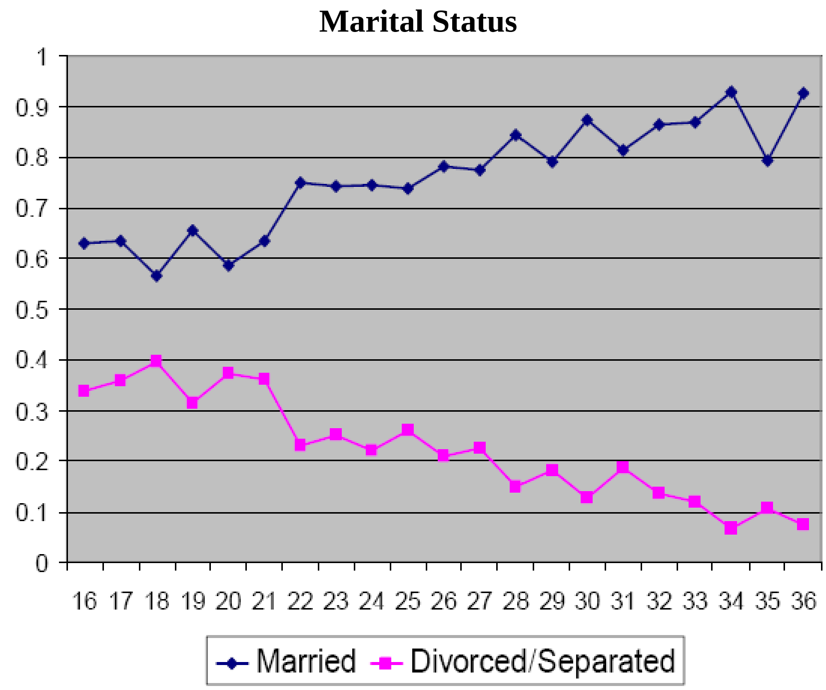
Source: National Longitudinal Survey of Youth, 2004. AFQT percentile scores are normalized by the age at which the woman took the test.

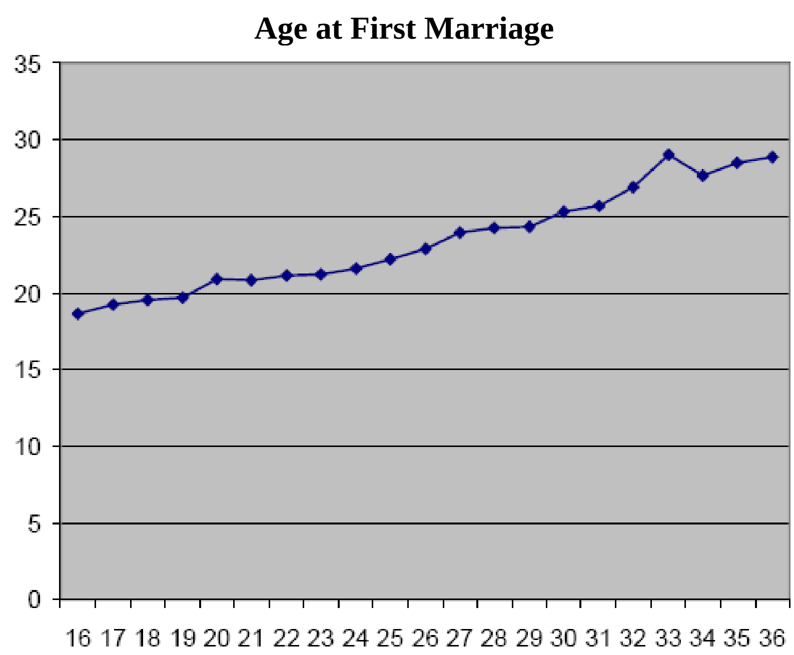
Figure 7: Percent of Women with a Child and Number of Children, by Age and Final Education



Source: National Longitudinal Survey of Youth, 1979-2004.

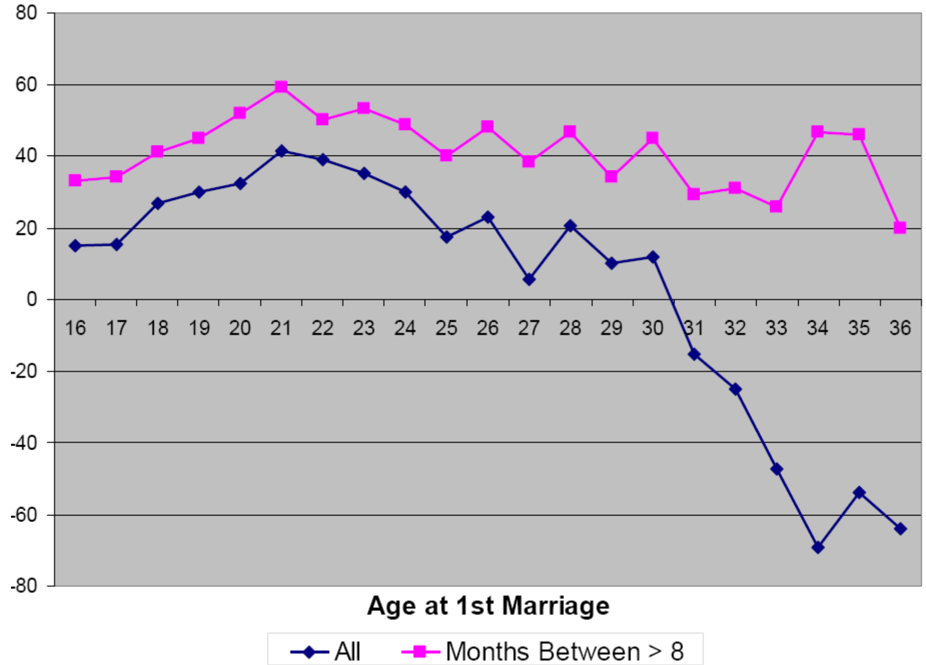
Figure 8: Marital Status and Age at First Marriage, by Age at First Birth





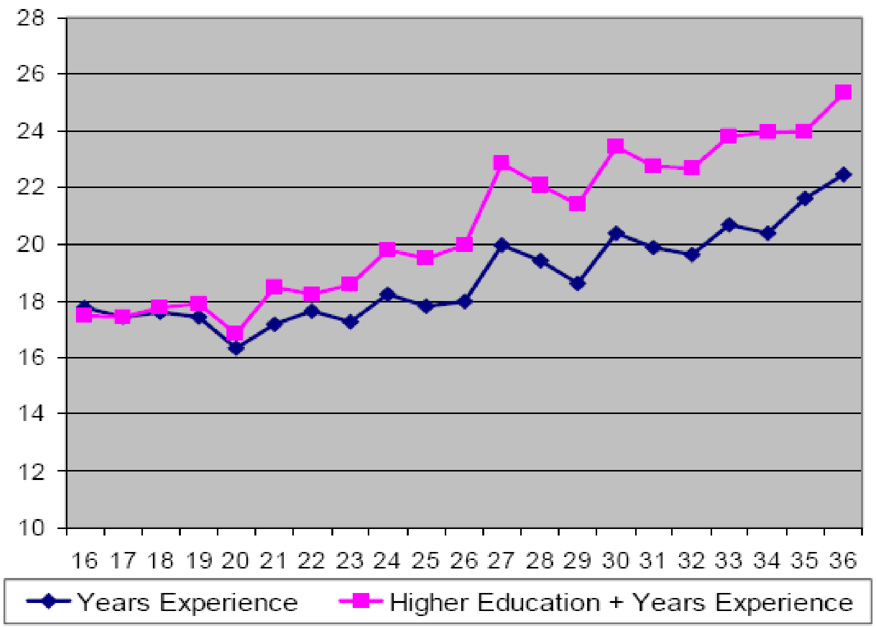
Source: National Longitudinal Survey of Youth, 1979-2004.

Figure 9: Months Between First Marriage and First Birth, by Age at First Marriage



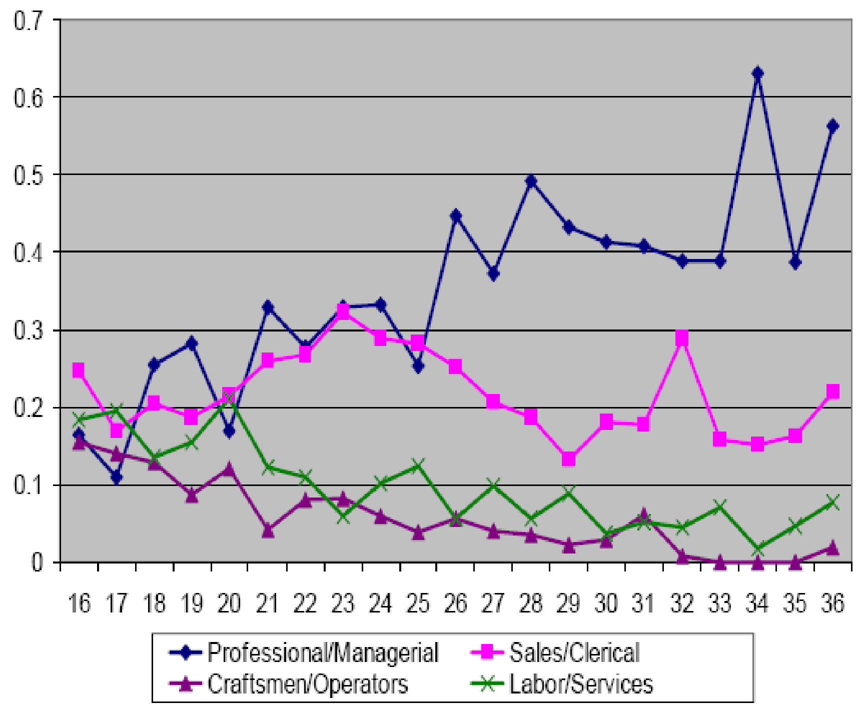
Source: National Longitudinal Survey of Youth, 2004. For Months Between > 8, the sample is restricted to those women who reported at least nine months between their first marriage and first birth.

Figure 10: Total Years of Experience, by Age at First Birth, 2004



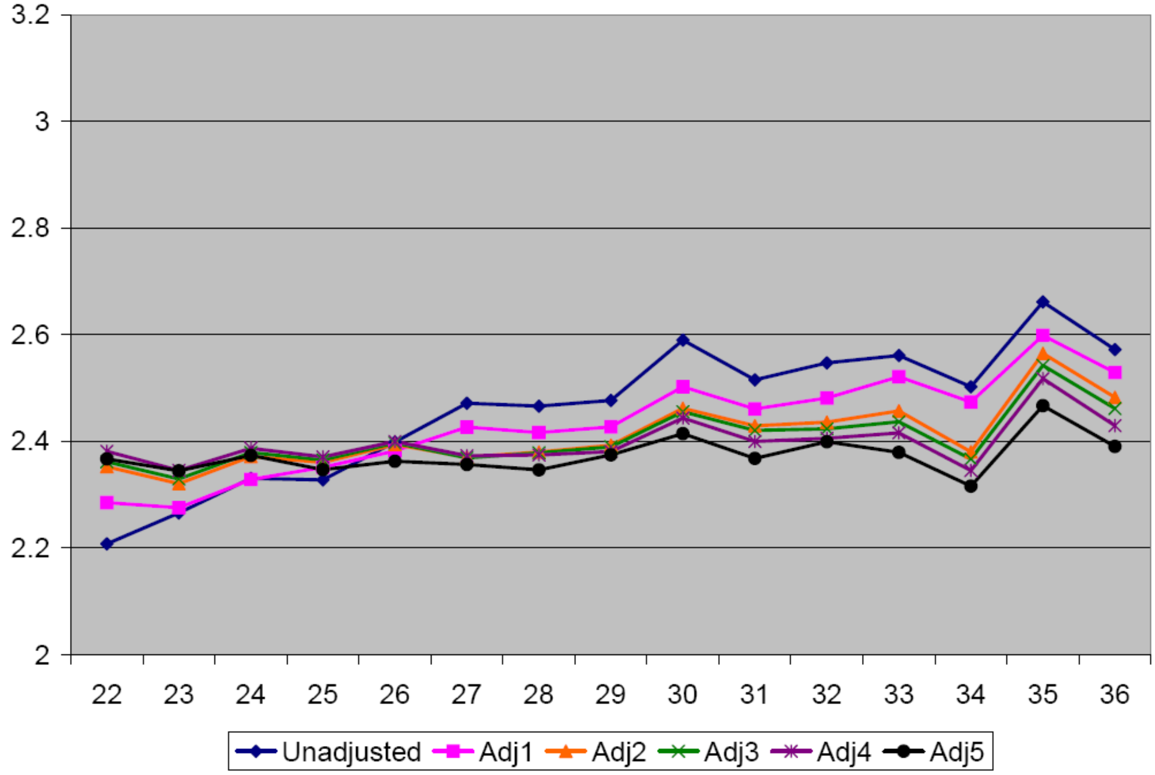
Source: National Longitudinal Survey of Youth, 1979-2004. Years of experience is measured as total hours worked between 1978 and 2003, divided by 2000. "Higher Education" is years of education completed above the high school level.

Figure 11: Occupation, by Age at First Birth



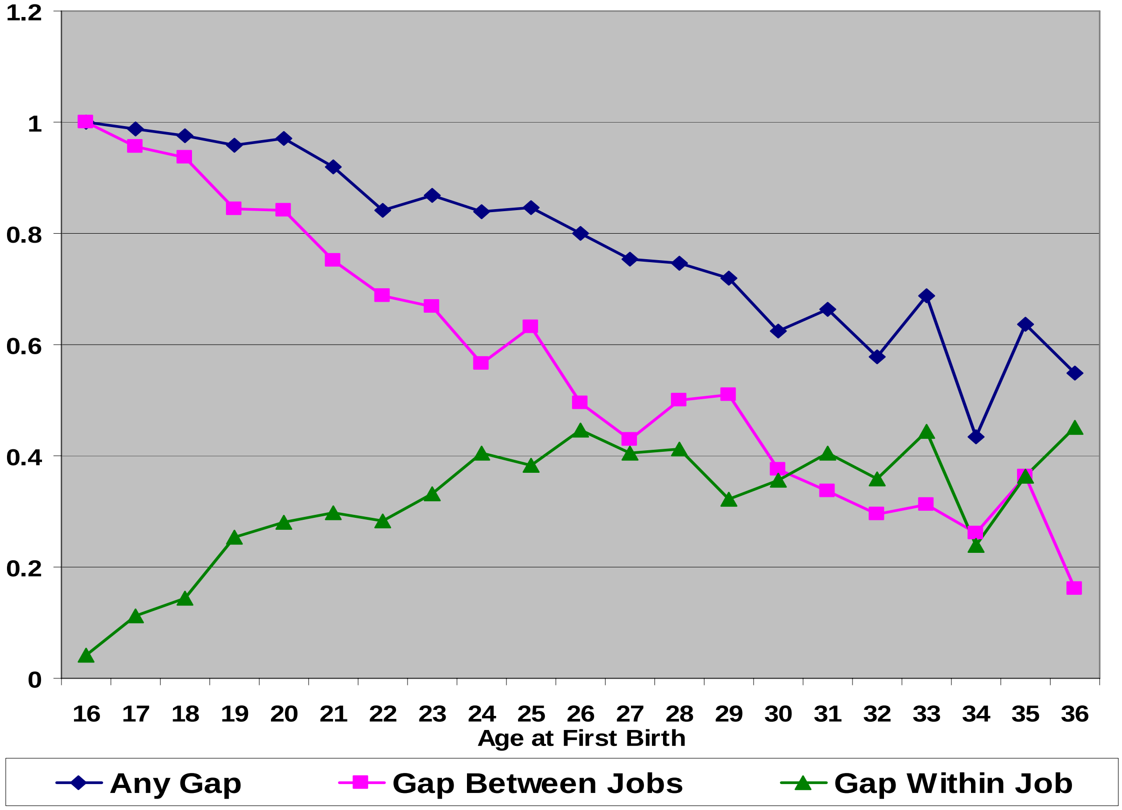
Source: National Longitudinal Survey of Youth, 2004.

Figure 12: Adjusted Log Hourly Wages by Age at First Birth



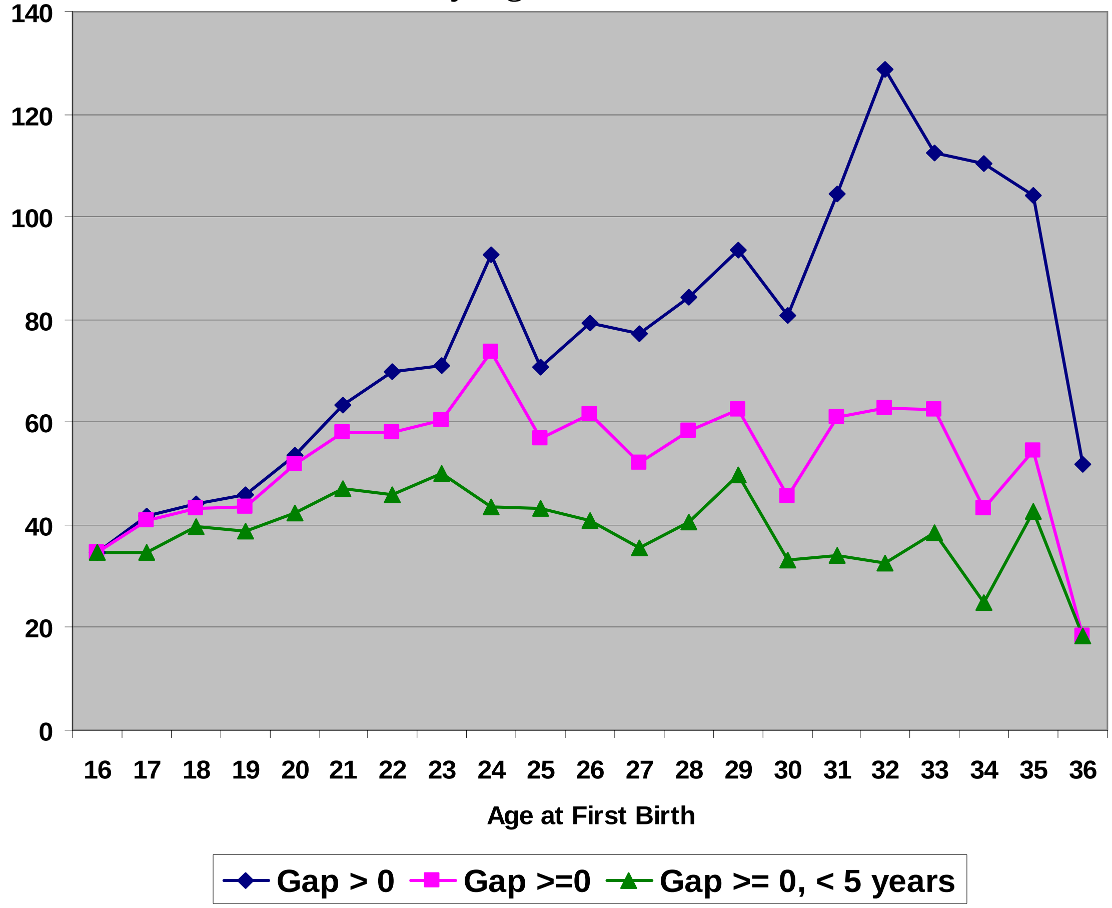
Source: National Longitudinal Survey of Youth, 1979-2004. Log hourly wages are adjusted for observable characteristics using random-effects regressions. All regressions are adjusted for age; Adj1 adds controls for AFQT score and race; Adj2 adds controls for education; Adj3 adds controls for age at first marriage; Adj4 adds controls for number of children and marital status; Adj5 adds controls for actual experience, experience squared, and occupation categories. The sample is limited to women with a first birth after age 21 who are 20 or older at the time of interview and with hourly wages between $1 and $200. The dependent variable is log hourly wages in year-2000 dollars, and experience is hours ever worked divided by 2000.

Figure 13: Percent of Women with Any Labor Force Interruption Around First Birth, by Age at First Birth



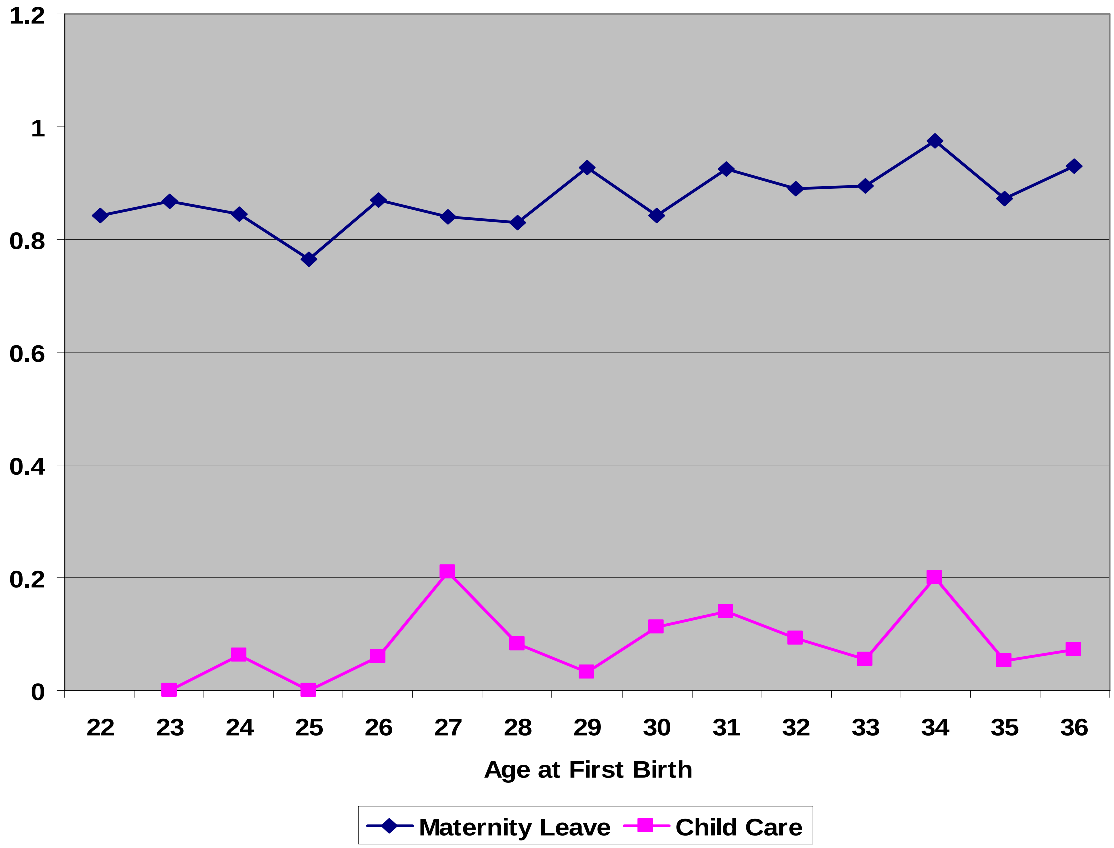
Source: National Longitudinal Survey of Youth, 1979-2004. Labor force interruptions are described as around the first birth if they occur in the interval beginning 40 weeks before the birth or 20 weeks after.

Figure 14: Average Length of First-Birth Employment Gap in Weeks, by Age at First Birth



Source: National Longitudinal Survey of Youth, 1979-2004. Labor force interruptions are described as around the first birth if they occur in the interval beginning 40 weeks before the birth or 20 weeks after.

Figure 15: Percent of Women with Maternity Leave Coverage or Employer-Proveded Child Care, by Age at First Birth



Source: National Longitudinal Survey of Youth, 1979-2004.

Table 1: Percentage Change in Wages after Birth of First Child, by Time Since Birth, Age at Birth, and Skill Level

		AFQT		
	(1) All	(2) Low	(3) Middle	(4) High
Group 1: First Birth Age 22-26				
Years 0-4	**-10.40**	**-6.19**	-4.12	**-18.69**
Years 5-9	**-21.47**	**-10.57**	**-16.56**	**-32.58**
Years 10-17	**-37.99**	**-23.57**	**-30.91**	**-51.86**
Years 18+	**-45.06**	**-29.97**	**-29.92**	**-61.89**
Group 2: First Birth Age 27-31				
Years 0-4	**-6.46**	-1.01	-3.03	**-13.23**
Years 5-9	**-25.75**	**-12.22**	**-23.93**	**-35.65**
Years 10-17	**-39.07**	**-25.43**	**-35.05**	**-51.43**
Years 18+	**-50.57**	**-43.88**	**-47.19**	**-58.47**
Group 3: First Birth Age 32+				
Years 0-4	**-7.95**	**-12.00**	-3.51	**-10.10**
Years 5-9	**-22.16**	**-21.89**	-9.63	**-30.80**
Years 10-17	**-38.67**	**-27.70**	**-35.58**	**-46.41**
Years 18+	-	-	-	-
Observations	2401	754	771	741
R-squared	0.4951	0.4394	0.4689	0.4970

Source: National Longitudinal Survey of Youth, 1979-2004. Results are based on fixed effects regressions with clustered standard errors. Percentage change should be interpreted as the percent difference in wages, compared to a woman of the same age who has not yet had a child. Numbers in bold are based on coefficients that are significant at the 5% level. Regression coefficients and standard errors are available in Appendix Table 1. The sample is limited to women with a first birth after age 21 who are 20 or older at the time of interview and with hourly wages between $1 and $200. The dependent variable is log of hourly wages in year-2000 dollars. The "Years" variables are dummies indicating years since first birth. Regressions also include a control for age.

Table 2: Effect of an Additional Year of Delayed Childbearing on Adjusted Log Wages, NLSY 1979-2004

Table 2a

Wages Adjusted For:	All 22-36	AFQT Low	AFQT Middle	AFQT High
Age	**0.030**	**0.014**	**0.021**	**0.032**
	(0.001)	(0.002)	(0.002)	(0.002)
+ AFQT, Race	**0.021**	**0.015**	**0.022**	**0.031**
	(0.001)	(0.002)	(0.002)	(0.002)
+ Education	**0.011**	**0.009**	**0.013**	**0.015**
	(0.001)	(0.002)	(0.002)	(0.002)
+ Age 1st Marriage	**0.009**	**0.006**	**0.012**	**0.011**
	(0.001)	(0.002)	(0.002)	(0.002)
+ #Kids, Marital Status	**0.005**	**0.004**	**0.008**	**0.006**
	(0.001)	(0.002)	(0.002)	(0.002)
+ Experience, Occupation	**0.003**	0.002	**0.004**	**0.005**
	(0.001)	(0.002)	(0.002)	(0.002)
Observations	2001	615	697	689

Table 2b

Wages Adjusted For:	All 22-36	AFQT Low	AFQT Middle	AFQT High
Age	**0.030**	**0.014**	**0.021**	**0.032**
	(0.001)	(0.002)	(0.002)	(0.002)
+ Experience, Occupation	**0.020**	**0.007**	**0.012**	**0.023**
	(0.001)	(0.002)	(0.002)	(0.002)
+ #Kids, Marital Status	**0.017**	**0.006**	**0.009**	**0.019**
	(0.001)	(0.002)	(0.002)	(0.002)
+ Age 1st Marriage	**0.011**	**0.003**	**0.005**	**0.007**
	(0.001)	(0.002)	(0.002)	(0.002)
+ Education	**0.004**	0.001	**0.004**	**0.004**
	(0.001)	(0.002)	(0.002)	(0.002)
+ AFQT, Race	**0.003**	0.002	**0.004**	**0.005**
	(0.001)	(0.002)	(0.002)	(0.002)
Observations	2001	615	697	689

Source: National Longitudinal Survey of Youth, 1979-2004. Results are the coefficients from a regression of adjusted log hourly wages regressed on age at first birth. The adjusted wages are the residuals from random-effects regressions of log hourly wages in year-2000 dollars on sets of controls, as indicated in the left column. The results in Table 2b are from regressions which add the independent variables in reverse order from Table 2a. A Hausman test was unable to reject the null hypothesis of equivalence of a fixed-effects and random-effects model. The sample is limited to women with a first birth after age 21 who are 20 or older at the time of interview and with hourly wages between $1 and $200. Women are classified into AFQT terciles by their percentile scores relative to women who took the test at the same age. Experience is hours ever worked divided by 2000.

Table 3: Effect of an Additional Year of Delayed Childbearing on Adjusted Log Wages, NLSY 2004

Table 3a

Wages Adjusted For:	All 22-36	AFQT Low	AFQT Middle	AFQT High
Age	**0.031**	0.005	**0.021**	**0.038**
	(0.005)	(0.010)	(0.009)	(0.009)
+ AFQT, Race	**0.020**	0.007	**0.021**	**0.036**
	(0.005)	(0.010)	(0.008)	(0.009)
+ Education	**0.014**	0.002	0.016	**0.026**
	(0.005)	(0.010)	(0.008)	(0.009)
+ Age 1st Marriage	**0.012**	0.005	0.012	**0.020**
	(0.005)	(0.010)	(0.008)	(0.009)
+ #Kids, Marital Status	0.008	0.001	0.009	0.014
	(0.005)	(0.010)	(0.008)	(0.009)
+ Experience, Occupation	0.006	0.001	0.007	0.008
	(0.005)	(0.010)	(0.008)	(0.008)
Observations	1075	315	359	388

Table 3b

Wages Adjusted For:	All 22-36	AFQT Low	AFQT Middle	AFQT High
Age	**0.031**	0.005	**0.021**	**0.038**
	(0.005)	(0.010)	(0.009)	(0.009)
+ Experience, Occupation	**0.018**	-0.001	0.012	**0.022**
	(0.005)	(0.010)	(0.008)	(0.008)
+ #Kids, Marital Status	**0.018**	0.001	0.013	**0.020**
	(0.005)	(0.010)	(0.008)	(0.008)
+ Age 1st Marriage	**0.015**	0.005	0.008	**0.010**
	(0.005)	(0.010)	(0.008)	(0.009)
+ Education	0.008	0.001	0.006	0.006
	(0.005)	(0.010)	(0.008)	(0.008)
+ AFQT, Race	0.006	0.001	0.007	0.008
	(0.005)	(0.010)	(0.008)	(0.008)
Observations	1075	315	359	388

Source: National Longitudinal Survey of Youth, 1979-2004. Results are the coefficients from a regression of adjusted log hourly wages on age at first birth. The adjusted wages are the residuals from OLS regressions of 2003 log hourly wages on sets of controls, as indicated in the left column. The results in Table 3b are from regressions which add the independent variables in reverse order from Table 3a. The sample is limited to women with a first birth after age 21 with hourly wages between $1 and $200. Women are classified into AFQT terciles by their percentile scores relative to women who took the test at the same age. Experience is hours ever worked divided by 2000.

Appendix Table 1: Coefficients Estimating Effect of First Child on Log Wages, by Time Since Birth, Age at Birth, and Skill Level

		AFQT		
	(1) All	(2) Low	(3) Middle	(4) High
Group 1: First Birth Age 22-26				
Years 0-4	**-0.1098**	**-0.0639**	-0.0421	**-0.2069**
	(0.0174)	(0.0278)	(0.0299)	(0.0345)
Years 5-9	**-0.2417**	**-0.1117**	**-0.1811**	**-0.3943**
	(0.0243)	(0.0425)	(0.0415)	(0.0451)
Years 10-17	**-0.4778**	**-0.2688**	**-0.3698**	**-0.7311**
	(0.0323)	(0.0574)	(0.0550)	(0.0592)
Years 18+	**-0.5990**	**-0.3562**	**-0.3555**	**-0.9646**
	(0.0431)	(0.0777)	(0.0727)	(0.0792)
Group 2: First Birth Age 27-31				
Years 0-4	**-0.0668**	-0.0102	-0.0308	**-0.1419**
	(0.0194)	(0.0385)	(0.0336)	(0.0308)
Years 5-9	**-0.2977**	**-0.1303**	**-0.2735**	**-0.4409**
	(0.0290)	(0.0587)	(0.0499)	(0.0466)
Years 10-17	**-0.4954**	**-0.2934**	**-0.4316**	**-0.7221**
	(0.0370)	(0.0765)	(0.0639)	(0.0581)
Years 18+	**-0.7046**	**-0.5777**	**-0.6384**	**-0.8787**
	(0.0698)	(0.1636)	(0.1168)	(0.0938)
Group 3: First Birth Age 32+				
Years 0-4	**-0.0828**	**-0.1278**	-0.0357	**-0.1065**
	(0.0308)	(0.0645)	(0.0476)	(0.0485)
Years 5-9	**-0.2505**	**-0.247**	-0.1013	**-0.3681**
	(0.0402)	(0.0719)	(0.0649)	(0.0662)
Years 10-17	**-0.4889**	**-0.3244**	**-0.4397**	**-0.6239**
	(0.0789)	(0.1073)	(0.0883)	(0.1913)
Years 18+	-	-	-	-
Observations	2401	754	771	741
R-squared	0.4951	0.4394	0.4689	0.4970

The coefficients from this table are used to create the percentage change results in Table 1.
Source: National Longitudinal Survey of Youth, 1979-2004. Results are coefficients from a fixed effects regression, with clustered standard errors. Coefficients in bold are significant at the 5% level. The sample is limited to women with a first birth after age 21 who are 20 or older at the time of interview and with hourly wages between $1 and $200. The dependent variable is log of hourly wages in year-2000 dollars. The "Years" variable refers to years since first birth. Regressions also include a control for age.

Appendix Table 2: Effect of Lagged Log Wage on Probability of a Missing Wage Observation After the Birth of the First Child: Marginal Effects From a Probit Regression

Dependent Variable = 1 if Missing Wage Observation		**AFQT**		
	All	**Low**	**Middle**	**High**
Group 1: First Birth Age 22-26				
dF/dx	**-0.1066**	**-0.1320**	**-0.1069**	**-0.0888**
z	-6.77	-4.71	-3.71	-3.30
Observations	2,214	796	746	546
Group 2: First Birth Age 27-31				
dF/dx	**-0.1191**	**-0.1454**	**-0.1526**	**-0.0727**
z	-5.92	-3.14	-3.93	-2.39
Observations	1,314	274	416	574
Group 3: First Birth Age 32+				
dF/dx	**-0.0748**	-0.1015	-0.0432	**-0.0784**
z	-2.78	-1.72	-0.60	-2.06
Observations	544	124	178	228

Source: National Longitudinal Survey of Youth, 1979-2004. Each entry represents the marginal effect (dF/dx) from a separate probit regression, where the dependent variable is equal to one if the wage observation is missing, and the independent variable is the log of the last reported real wage. Regressions also include a control for additional children born since the first birth, and the number of years since the last reported wage observation. The sample is limited to women with a first birth after age 21 who are 20 or older at the time of interview, who report having a first birth within the last two years. Coefficients in bold are significant at the 5% level.