The Effects of Motherhood Timing on Career Path

July 2005

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Abstract

This paper estimates the effects of motherhood timing on female career path, using national panel data from the NLSY79, and biological fertility shocks as instrumental variables for the age at which a woman bears her first child. Motherhood delay leads to a substantial increase in career earnings of 10% per year of delay, a smaller increase in wage rates of 3%, and an increase in career hours worked of 5%. The postponement premium is largest for college-educated women, and those in professional and managerial occupations, which supports a human capital story for the timing effect. Family leave laws do not significantly influence the premium. Panel estimation reveals evidence of both fixed wage penalties and lower returns to experience for mothers. Conversely, using measured aptitude level as an instrumental variable for expected future earnings, we show that higher expected career earnings lead women to postpone childbearing.

* I thank Susan Athey, B. Douglas Bernheim, John Pencavel and Antonio Rangel for their excellent advising, and Katie Carman, Victor Fuchs, Edward Lazear, Laura Lindsey, David Matsa, Carmit Segal, Marianne Simonsen, Scott Stern, Ed Vytlacil and Geoff Warner for helpful comments. The Koret Foundation and the Stanford Institute for Research on Women and Gender provided generous financial support.

1. Introduction

Delayed motherhood is associated with higher female career achievement in both cross-sectional and time-series comparisons. Fertility delay has been increasing concurrently with female education, labor force participation, and earnings in Europe since 1960 (Gustafsson, 2001), and in the United States since the post-war baby boom (Chen and Morgan, 1991; Caucutt et al., 2002). Hofferth (1984) noted the cross-sectional correlation almost two decades ago in data from the 1976 Panel Study of Income Dynamics (PSID). Women who bore their first child after age 30 enjoyed higher wage rates and accumulated more wealth by age 60 than earlier child-bearers and childless women.

An analogous relationship is evident for the more recent cohort of women in the National Longitudinal Survey of Youth 1979 sample (NLSY79). The first five figures of this paper chart career profiles in median earnings, hours worked, and wage rates for women grouped by the age at which they first gave birth (A1B). Figure 1 shows the progression of raw log-earnings over time, while Figure 2 is adjusted for differences in education, race, ability (Armed Forces Qualification Test score) and birth year cohort. Since hours are substantially different across the groups, as shown in Figure 3, log-wage rates may be a more appropriate focus; these are depicted in Figures 4 and 5. Even in Figure 5, with adjusted values, mothers with age at first birth (A1B) in the range 30 to 34 outperform childless women and the two groups of earlier mothers (A1B 20 to 24 and A1B 25 to 30) by the age of 35. Differences in wage rates and wage growth rates are apparent in Figure 5 even prior to childbearing, implying a role for unobserved variables such as ambition or productivity to explain the correlation between delayed motherhood and earnings.

Without further evidence, the correlation can be attributed to motherhood delay affecting earnings, anticipated earnings affecting motherhood delay, or some outside factor influencing both. This paper will attempt to isolate the first two effects. In defining the effect of delay, I begin with the following thought experiment. A particular woman, with a given level of ability, career motivation, labor force attachment and education, wants to have a single child, and spend a fixed amount of time out of the labor market caring for her child. The first question is how much does motherhood timing itself - having a child at age 24 versus age 27 - influence subsequent wage rates and lifetime earning potential, holding all other factors fixed? Turning to the converse question, I ask: how does a woman's earning potential affect her choice of motherhood timing? Consider two identical women, who differ only in their potential wage profiles. Does the woman who anticipates higher earnings prefer to delay motherhood longer? Both aspects are important for researchers interested in women's labor supply and in the effects of motherhood and career interruptions on their earnings.

From a policy perspective, motherhood delay matters because of its potential effects on maternal and infant health, population size, and sexual equity. The social and private costs from "late" motherhood have garnered much attention in recent years. Deferred motherhood contributes to a reduction in population growth, a major concern in Europe, where total fertility is below replacement level. Gustafsson (2001) points out that average age at first birth has reached an "all time high" in many European countries, accounting for much of the fertility decline, and argues for increased government action. Older mothers may also face additional health costs, as detailed in "Delayed Childbearing" (*Atlantic Monthly,* 1995). For example, childbearing after age 35 or 40 can be difficult and costly to achieve, and has been linked to a greater incidence of birth defects and congenital anomalies. A recent study (Alonzo, 2002) connects motherhood after age 35 with heightened risks of heart attack, congestive heart failure, high blood pressure, diabetes and dental and vision problems.[1] Policymakers concerned with fertility deferment can use the estimate in this paper to understand how much of the delay arises from career concerns, and to design appropriate incentives, potentially targeted at different populations of women.

Motherhood may be the remaining obstacle to women's achievement of economic equality with men (Fuchs, 1988), and deferred motherhood may be a mean of reducing that inequality. At the same time, any financial rewards to motherhood delay are themselves a central component of the work-family conflict. Hewlett (2002) argues that young women have been misled regarding the cost of fertility deferment, and as a result, come to regret the family sacrifices they make for professional success. In that

[1] On the other hand, later mothers experienced reduced risks of stroke and bladder infection.

context, the estimates of this paper can be a gauge of the difficulties women face in balancing work and family.[2] This paper will not discuss the personal costs of delay or solve for a private or social *optimal* timing of motherhood. Instead, it focuses on measuring the potential career rewards from delay, and the fertility response to potential earnings.

The core results are that fertility delay influences career path, and career potential affects fertility timing. A year of delayed fertility leads to a 10% increase in career earnings, a 5% increase in career work experience, and a 3% increase in career average wage rate. The effects are not the same for all women, and women with college degrees, and those in professional and managerial occupations receive the greatest returns. Surprisingly, family leave laws are not shown to alleviate the tradeoff. Panel estimation reveals evidence of both fixed wage penalties and lower returns to experience for mothers, and suggests that these costs are lower for older mothers. Finally, using measured aptitude level as an instrumental variable for expected future earnings, I show that higher expected career earnings lead mothers to postpone childbearing, where a doubling of earnings is associated with about a year and a half of delay.

2. Related Literature

This paper presents new evidence concerning how labor markets respond to motherhood by exploring the effects of timing on the career costs of childbearing. The wage differential between mothers and childless women has been termed the "family gap" and studied extensively by economists and sociologists (e.g. Cramer, 1980; Browning, 1992; Joshi et al., 1998; Waldfogel, 1998; Dankmeyer, 1996; Budig and England, 2001). Korenman and Neumark (1992) describes how unobserved heterogeneity and endogenous fertility present obstacles to drawing causal inference from cross-sectional ordinary least squares (OLS) estimates. They estimate the family gap using OLS, first-differences and instrumental

[2] Family and children's advocates (Crittenden, 2001; Atkinson, 2003) have argued that the tradeoffs are unfair, and signal a need for extended government intervention in supporting working mothers – through mandated paid maternity leave; an extension of current unpaid leave to all workers; or through universal access to quality and affordable childcare services. Fairness is outside of the scope of this paper.

variables (IV) using National Longitudinal Surveys 1968 (NLS68) data. Their instrumental variables are proxies for family background and for "beliefs and expectations."[3] The major weakness of these instruments is that, while they may predict motherhood, there are few *a priori* reasons to exclude them from the wage equation. A recent study by Simonsen and Skipper (2003) measures the family gap using propensity score matching, a procedure that is more flexible regarding functional form than OLS, but which also depends on motherhood being conditionally independent of wages or on finding a valid exclusion restriction.[4] Biological fertility events associated with twins (Bronars and Grogger, 1994) and sex ratios (Angrist and Evans, 1998) have been exploited as instrumental variables for number of children in the wage equation, but they cannot be used to identify the effects of delaying or avoiding childbearing. Section 7.2 of this paper estimates the effects of motherhood on the wages profile using biological fertility shocks as instrumental variables.

An alternative approach to measuring motherhood costs is to focus on career interruptions and the timing of work experience. Light and Ureta (1995) find that experience timing is related to wages, and can account for 12% of the male-female wage gap. Mincer and Polacheck (1974) and Mincer and Ofek (1982) describe the role of human capital investment during work years, and find evidence of its depreciation during career breaks for women in the NLS68. Depreciation rises with education level, and is higher for women with more years of experience at the time of their interruption.[5] Using data from the more recent NLSY79, Baum (2002) finds evidence of depreciation, but only among women who switched employers. Albrecht et al. (1999) reproduces the result with Swedish data, and argues for the importance of asymmetric information and signaling in the costs of career interruptions. They find that interruption type, and not simply duration, determines the wage penalty, and that family leave is more harmful to men than to women. While the findings of "depreciation" are generally consistent, an alternative interpretation

[3] The background variables are: parental education, parental educational goal for respondent, number of siblings, mother's labor force participation, and family disruption. And the belief variables are: indicators for strong agreement and disagreement with a statement about women working against her husband's wishes, ideal age at marriage (at age 14), expected number of children (in 1980), and educational goals (in 1970) and expectations (in 1970).
[4] The authors do both. In conducting residual based matching, they use number of siblings as the predictor of motherhood that they exclude from the wage equation.
[5] This may be caused by women who expect larger depreciation costs choosing to delay motherhood.

is that women choose to have children at times when they expect lower wages. Gronau (1988) notes this problem, and discusses the challenge in identifying causal effects between wage differentials and career interruptions. This paper identifies the effects of motherhood delay, which are related, but not identical, to the generic effects of experience timing. As such, the results may not generalize to other types of career interruptions.[6]

The impact of motherhood timing itself has also received attention in several studies using similar estimation methods as those used to measure the family gap. The finding in Hofferth (1983) depends on the assumption that motherhood timing is exogenous, conditional on covariates, in her cross-sectional data. Taniguchi (1999) uses (NLS68) longitudinal data and controls for fixed individual differences in earnings. Since age at first birth is static over time, her variable of interest is the product of A1B category indicators and number of children: she finds that motherhood penalties are largest for women who begin childbearing between 20 and 27, are lower for teen mothers, and are negligible for older mothers. While the fixed effects absorb variation that is expressed through vertical shifts of the age-wage profile, the technique fails to account for other differences in wage profiles or for responsiveness of motherhood timing to career outcomes.

As in the family gap literature, the IV studies use socioeconomic background and "beliefs" as instrumental variables (Blackburn et al., 1993; Chandler et al., 1994; and Amuedo-Dorantes and Kimmel, 2003).[7] The instruments are susceptible to similar criticism here, as there is little reason to assume that they influence fertility timing but not wage rates. Further, since women are generally aware of these factors early in life, they may respond to fertility timing expectations through career choice or unobservable investments that in turn influence wages. Amuedo-Dorantes and Kimmel (2003) finds that delaying motherhood past age 30 eliminates the family gap, but only for college-educated women. A

[6] In fact, when I control for the timing of work experience, the OLS coefficient for A1B (for earnings or for wages) is reduced by half. The diminished effect is still statistically significant at the 1% level, suggesting that motherhood does have a distinctive impact or that more complex experience timing interactions are important.
[7] Blackburn et al. (1993) uses number of siblings, parental education, mother's labor force participation, and family disruption. Chandler et al. (1994) uses religion and number of siblings. Amuedo-Dorantes and Kimmel (2003) uses parental education and family disruption.

similar interaction between education and motherhood delay is presented in Section 7.1. The other two studies deemphasize their IV results. Blackburn et al. (1993) finds no effect of motherhood delay on earnings using IV. Chandler et al. (1994) reports OLS findings of a 1% wage benefit per year of delay, and describes the IV estimates as similar, but sensitive to model specification.

This paper contributes an approach to identification using biological fertility shocks as instrumental variables for fertility timing in the career outcome equations. The strategy is similar in spirit to Hotz et al. (1999), which uses miscarriage (spontaneous abortion) as a natural experiment to consider the impact of teenage motherhood. While they find no benefit from motherhood delay, their results need not extend to later delays, as older mothers face substantially different constraints in career and family; older mothers tend to have higher earnings potential and their motherhood timing conflicts are more likely to involve work than school. The instrumental variables in this paper are: (1) whether first pregnancy ended in miscarriage, (2) whether conception of first child occurred while using contraception, and (3) duration, in years, of the conception attempt prior to first birth. These factors each shift a woman's actual age at first birth away from her preferred age. They are plausibly exogenous to earnings potential, and largely unanticipated.[8]

Turning to the converse relationship, previous work has associated fertility timing with male and female wages in time-series (Butz and Ward, 1979) and cross-sectional data (Heckman and Walker, 1990). Happel et al. (1984) finds that women in higher-paying occupations choose longer intervals from first marriage to motherhood, and Caucutt et al. (2002) associates higher female earnings with delayed motherhood. These correlations are similar to the ones noted in Hofferth (1984), only viewed from a different perspective. In order to isolate the causal effect of anticipated earnings on motherhood delay, I use aptitude as measured by the Armed Forces Qualification Test (AFQT) score as an instrumental variable for expected earnings potential. The identifying assumption is that controlling for education level and other covariates and apart from career considerations, more intelligent women do not prefer longer delays before motherhood.

[8] The validity of the instruments is discussed further in Section 4.

3. Conceptual Framework

In this section, I present benchmark cases to illustrate some potential labor market effects of motherhood. This exercise motivates empirical choices of outcome variables and an investigation of the underlying impact of motherhood on wages. To isolate the labor market effects of motherhood delay, I focus on scenarios with motherhood career interruptions of fixed duration. Consider the hypothetical woman from Section 1, who has a single child, works full-time until motherhood, takes a year out of the labor force to care for her newborn, and returns to full-time employment. The effects of motherhood delay on her career outcomes are considered under five different scenarios: (1) discounting and no wage growth, (2) secular wage growth greater than discounting, (3) foregone human capital accumulation during labor market absence, (4) a fixed penalty for motherhood or interruptions, and (5) lower returns to work experience for mothers. While some of these costs may generalize to the timing of any career interruption, the examples in this section are for motherhood, keeping in line with the empirical work which exploits exogenous variation in fertility timing, but not other aspects of experience timing.

In Case 1, the woman discounts future income and works in a setting with no wage growth. The labor market costs of motherhood arise from foregone wages during her absence. The present value of the foregone wages is lower for later interruptions because of discounting, leading to financial benefits to delay. In Case 2, the woman experiences wage growth over time greater than the discount rate, causing a reversal in the financial effect of motherhood delay. Early motherhood is relatively less costly since the price of home time is rising. Steeper age-wage profiles lead to larger financial incentives for early motherhood. In neither of these cases do wages depend on experience or do terminal wage rates vary with fertility timing.

In the next three cases, wages rise with age because of increased worker productivity from on-the-job training and investment in human capital. The cost of a work interruption is the sum of the foregone wages and the foregone human capital accumulation:

$$\text{Foregone Wages} + \text{Foregone Human Capital} = \int_{A1B}^{A1B+Break} w(t)e^{-rt}dt + \int_{A1B+Break}^{T} [w(t) - w(t, A1B)]e^{-rt}dt$$

where Break is the duration of the labor force interruption, T is age at retirement, *w(t)* are wages for a childless woman, and *w(t,A1B)* are wages for a mother who bore her child at age A1B. Having discussed discounting above, I set the discount factor *r* to zero for the remaining cases. The first term increases in A1B, capturing the effect of wage growth in raising foregone wages for later interruptions. The second term deceases in A1B, as the foregone human capital investment is accrued over fewer years for later interruptions.

The equation describes Case 3, illustrated in Figure 6, where wages increase with experience, and there are no additional costs to motherhood. The highest line in the figure represents the uninterrupted age-wage profile. The effects of motherhood are shown, for A1B=23 and A1B=30, as: (1) zero earnings during the interruption and (2) a return to pre-motherhood wage level after the break. Since the profile depends solely on experience, the effect of a break is simply a horizontal shifting of the profile for the duration of the break. Irrespective of motherhood timing, the terminal profile is the same, and terminal wage rates do not depend on fertility timing. Career earnings do not vary with timing either. They are equal to the area under the wage rate line, and an interruption is simply a slice and shift of the shape. The location of the slice does not change the total area, and the total cost of any interruption is the value of earnings during the last period (equal in duration to the interruption) on the uninterrupted profile. This equivalence does not depend on a particular functional form for the relationship between wages and experience.[9]

Case 4 introduces a fixed cost, which can be thought of as a motherhood penalty (lower productivity or employer response to a negative signal) or depreciation of human capital during career interruptions. A fixed and lasting motherhood penalty is represented in Figure 7 by a vertical downward shift in the mother's wage profile. Although terminal wage rates do not vary with motherhood delay,

[9] It does depend on the assumption that the duration of the break is independent of the timing of motherhood, and of no discounting. Also, if women work part-time after childbearing, the exact equivalence only holds in the case of linear returns to experience.

lifetime earnings increase, since later mothers work more years on the higher (un-depreciated) profile. Of course, human capital depreciation may take a form other than a fixed linear drop. Depreciation can be increasing in experience, if the loss of capital is proportional to the amount accumulated, or decreasing, if women with more seniority are better able to protect their human capital assets. In these situations, terminal wage rates will also vary with motherhood timing.

A final generalization of the baseline example (Case 5) allows for a reduction in returns to work experience following the birth of a child. This is depicted in Figure 8, where the age-wage profile is flatter following age at motherhood. Several factors can account for the flattening. Mothers may experience reduced opportunities for training and promotion, and find themselves relegated to a "mommy track."[10] Alternatively, older women or mothers may be less adept at developing new skills, and consequently earn smaller returns on their human capital investment. When early career contests conflict with early motherhood, women may benefit professionally from delayed motherhood.[11] In these situations, women who defer motherhood will receive higher earnings and higher terminal wages.

The benchmark cases considered in this section provide a useful conceptual framework, and highlight the potential channels through which timing of motherhood affects earnings and wages. Motherhood delay improves career earnings if there is a fixed cost of motherhood, a reduction in returns to experience, or discounting greater than secular wage growth. Delay leads to higher terminal wage rates only if depreciation costs are decreasing with experience, or if mothers experience a slope decrease. For this reason, I explore terminal wage rates (approximated by wages at age 34), in addition to career outcomes. Using panel data regressions, I also examine direct evidence of human capital depreciation and slope change following first birth. Section 5 relaxes certain simplifying assumptions by allowing multiple children, interruptions of varying duration, and part-time work. These labor supply effects of motherhood

[10] For an anecdotal account from the legal profession, see Mumford's account of having a child while an associate at a major New York law firm in *American Lawyer* (2003).
[11] This may be especially relevant for women working in academia, law, medicine, and accounting. For some discussion of in the context of academic careers, see Mason and Goulden (2002).

timing are accommodated by providing two specifications: one that controls for work experience, and one in which the total impact of delay is defined to include variation in experience.[12]

4. An Instrumental Variables Approach

Although delayed motherhood is positively correlated with female earnings, correlations alone do not to establish causality. The first step in isolating the effect of timing on earnings is to control for observable factors associated with both, which can be accomplished through a variety of estimation techniques that rely on the assumption that A1B is exogenous, conditional on covariates. An ordinary least squares (OLS) approach is to estimate the following equation:

$$\mathrm{Ln}(\mathrm{Y}_i) = \beta_0 + \beta_{A1B}\mathrm{A1B}_i + \beta_X \mathrm{X}_i + \varepsilon_i \qquad (1)$$

where Y_i represents earnings, wage rate or hours worked, $\mathrm{A1B}_i$ is age at first birth, X_i are the control variables, and ε_i is the error term. The parameter of primary interest is β_{A1B}, the effect of fertility delay on the outcome variable. OLS and the other methods produce biased estimates if A1B is correlated with the error term ε_i.

In practice, age at first birth will be correlated with the error term if unobserved factors that determine wages are correlated with motherhood timing. An example is preferences for family over career that both cause a woman to invest less effort or be less productive at work and to begin childbearing sooner. Conversely, women who are especially career oriented may tend to defer childbearing. Social group factors that affect both family and career choices, and are not captured by the usual controls for socioeconomic background and education, can also produce a spurious correlation. Finally, women with higher earnings potential and greater financial costs from motherhood may choose delay to reduce the penalty. Figures 1 to 5 of this paper imply that earnings differentials precede motherhood, even after controlling for observable factors. If later mothers face higher potential earnings

[12] In this case, the thought experiment needs to be altered to allow for the duration of the break to vary with motherhood timing. Hours worked result from some optimization that depends on age at first birth, and fertility timing leads to differences in hours and wages. The “total effect of delay” can be defined to include or to exclude the “hours effect.”

profiles than earlier mothers, as suggested by the figures, OLS will overstate the earnings benefits of deferred motherhood. Alternatively, if women with higher earnings potential are the ones who have children sooner, OLS will underestimate the effect of delay.[13]

This paper uses instrumental variables (IV) to identify β_{A1B}, exploiting biological factors that shift fertility timing exogenously to earnings. This method produces consistent estimates of the causal effect if the instruments are valid; they must be reasonably powerful predictors of fertility timing, and must not be correlated with the unexplained part of the outcome equations.

The instrumental variables for fertility timing are:

- an indicator for first pregnancy ending in miscarriage or stillbirth
- an indicator for first pregnancy occurring while mother was using contraception, and
- the lag in years from first attempt to conceive to first birth.

These variables capture random and unanticipated factors that drive a wedge between actual timing of motherhood and the woman's desired or optimal timing. The underlying model is that actual fertility timing is the result of the woman's desired timing and a series of random shocks, including the instruments, represented by the first stage regression:

$$\mathrm{A1B}_i = \gamma_0 + \gamma_M \mathrm{Miscarriage}_i + \gamma_C \mathrm{Contraception}_i + \gamma_L \mathrm{YearsToBirth}_i + \gamma_X \mathrm{X}_i + \varepsilon_i \qquad (2)$$

where X_i contains the controls from the wage equation and additional dummy variables for substance use during pregnancy (alcohol, smoking, marijuana and cocaine) and for having reported contraceptive use prior to first birth.

If the instrumental variables are correlated with age at first pregnancy attempt, they may not be exogenous to wages, a possible concern for miscarriage and time to conception, since female fecundity is known to decline with age. However, the drop is highly nonlinear, and is most apparent for pregnancies

[13] A story for this correlation is that women who are more successful and driven in their careers are also more "successful and driven" in their personal lives, marrying and bearing children sooner.

after age 33, the highest age at first birth included in the sample.[14] Another concern is that miscarriage may be associated with risky behavior. However, medical evidence does not support a strong impact of behavioral factors on miscarriage risk. Rather, over 85% of miscarriages occur within the first trimester of pregnancy and over 90% are caused by genetic defects or other anomalies that prevent the fetus from developing properly (Merck 1999). Miscarriage has been associated with some extreme behaviors such heavy alcohol use or drug addiction.[15] Fortunately, these factors are observed and controls are included in the regressions. While miscarriage may directly affect a woman's psychological state and work productivity, this is more likely for recurrent miscarriage, which affects less than 1% of women (Regan 2001).

Another concern arises regarding the "accidental pregnancy" variable (whether first pregnancy occurred while using contraception); women might differ in their contraceptive practices in ways that are correlated with earnings. In addition to including an indicator for contraceptive use prior to first birth, I also show in Section 7.2 that the effect of "accidental pregnancy" is not solely a result of differences in *type* of contraception used. Potential bias remains in correlations between a woman's consistency in her use of contraception or in her time to conception and her unobserved productivity at work. The advantage of multiple instruments for a single endogenous variable is that the system is over-identified and the restrictions are tested empirically.

Finally, the IV technique is easiest to interpret with homogenous treatment effects. The effects of motherhood delay are likely heterogeneous, complicating our interpretation of the β_{A1B} coefficient. In particular, IV estimates the local average treatment effect (Imbens and Angrist, 1994), weighted across women by their fertility timing response to the instrumental variables. Of course, some of the most

[14] One way to address the issue of declining fecundity is to estimate the effect of delay separately women who had their first child within different age ranges. Due to the small sample size, this exercise does not produce precise estimation results.
[15] See Regan (2001), pp. 120-122, for miscarriage risks associated with heavy alcohol use, drug use (cocaine, marijuana and cannabinoid compounds), and regular smoking during pregnancy.

important factors leading to heterogeneity are observable. Thus, I estimate the effects separately by education level, occupation type, employment sector and family leave regime.[16]

5. Empirical Analysis

5.1 The Effects of Motherhood Delay on Earnings, Hours Worked, and Wage Rates

The empirical analysis begins with direct estimation of the effect of motherhood delay on career outcomes. The unit of observation is an individual woman, and the baseline regression takes the form:

$$\mathrm{Ln}(\mathrm{Y}_i) = \beta_0 + \beta_{A1B}\mathrm{A1B}_i + \beta_X \mathrm{X}_i + \varepsilon_i \qquad (3)$$

where $\mathrm{A1B}_i$ is age at first birth, and the X_i controls are birth year cohort, education level, race (indicators for Black and Hispanic), and percentile score on the Armed Forces Qualification Test (AFQT), meant to capture general intelligence. This function is estimated using OLS, and then IV, using the instrumental variables listed above: miscarriage at first pregnancy, "accidental" first pregnancy occurring while using contraception, and time from first conception attempt to first birth. The dependent variables are selected to evaluate separately the effects of fertility timing on career earnings, career labor force experience, career wage rate, and terminal wage rate, as motivated in Section 3. Due to data constraints, "career" earnings are aggregated over the age range from 21 to 34, and "terminal" wages are at age 34, which allows the same age range to be used for women from different birth year cohorts.

Heterogeneity is explored by adding interaction terms between age at first birth and education level, occupation type, employment sector, spouse earnings, and family leave regime. In these specifications, the additional instrumental variables are the interactions of the original instruments with the exogenous variables of interest. My hypothesis regarding family leave is that it can reduce the returns to deferred motherhood by alleviating the career-motherhood conflict. If the motherhood penalty is concentrated among women who are not able to return to the pre-interruption jobs (as in Baum 2002),

[16] To explore heterogeneity in the effects of the instruments (rather than in the "treatment effect" itself), I repeated the first stage regressions on various sub-populations, such as A1B category and race. While the instruments were jointly significant for each of the groups, the coefficients varied somewhat, suggesting, for example, that failed contraception had a smaller effect on A1B for Hispanic and Black women.

then job protection should reduce the costs of interruption for all women. It should have the largest effect on the most vulnerable workers, young women with only brief established tenures. A negative A1B·FamilyLeave coefficient would indicate that delay is less important under Family Leave and support the hypothesis that mandated protection is an effective substitute for seniority in preserving women's career opportunities after motherhood.

5.2 The Effects of Motherhood on Wages and Wage Growth

In addition to estimating the overall career effects and interaction terms, I exploit the panel nature of the data to uncover characteristics of the response of wages to motherhood: specifically, for evidence of human capital depreciation or other fixed motherhood penalties, and for a reduction in returns to experience for mothers or a "mommy track."

The first panel regressions have $\text{Ln}(\text{Wage}_{it})$ as the dependent variable and include controls for individual fixed effects, age, motherhood status, and years since first birth. While more flexible representations are possible, the small sample size forces some parameterization to increase precision. The first specification includes only linear terms for wage growth, and the second adds quadratic terms:

$$\begin{aligned}\text{Ln}(\text{Wage}_{it}) = {} & \beta_0 + \beta_1 \text{Age}_{it} + \beta_2 \text{Age}^2_{it} + \beta_3 \text{Mother}_{it} + \beta_4 \text{Mother}_{it} \bullet (\text{Age}_{it} - \text{A1B}_i) + \\ & \beta_5 \text{Mother}_{it} \bullet (\text{Age}_{it} - \text{A1B}_i)^2 + \alpha_i + \varepsilon_{it}\end{aligned} \tag{4}$$

where $(\text{Age}_{it}\text{-A1B}_i)$ measures years since first birth, and α_i captures individual fixed effects. Instrumental variables estimation of this equation uses the predicted value from the first stage regression (Equation 2) with the original instruments ($\widehat{A1B}_i$) to create instruments for the motherhood indicator and interactions with years since first birth.[17] The new instruments are:

- $1(Age_{it} \geq \widehat{A1B_i})$, an indicator for Age being greater than predicted A1B,

[17] The original instrumental variables are fixed for a given woman, and cannot be used in a panel framework with individual fixed effects. Instead, they are interacted with the exogenous regressor *Age*, which is time-varying, to produce new instruments. The form of the new instruments was chosen to correspond to the endogenous variables.

- $1(Age_{it} \geq \widehat{A1B_i})\bullet(Age_{it} - \widehat{A1B_i})$, the indicator interacted with predicted years since first birth, and
- $1(Age_{it} \geq \widehat{A1B_i})\bullet(Age_{it} - \widehat{A1B_i})^2$.

In another specification, the quadratic terms are removed, and instead the effect of motherhood is allowed to vary with timing of first birth. The regression equation is:

$$\begin{gathered}\text{Ln}(\text{Wage}_{it}) = \beta_0 + \beta_1\text{Age}_{it} + \beta_2\text{Mother}_{it} + \\ \sum_{h=1}^{3}\beta_{3,h}\text{Mother}_{it}\bullet(\text{Age}_{it} - \text{A1B}_i)\bullet(\text{A1B}_i \in \text{FBCat}_h) + \alpha_i + \varepsilon_{it}\end{gathered} \qquad (5)$$

where the three FB Categories are: A1B between 20 and 24, A1B between 25 and 29, and A1B between 30 and 34. As the instruments are produced from a nonlinear function of the exogenous variables, traditional IV error estimates can be biased, and standard errors are computed by bootstrapping. Using the estimated parameters from the second stage, I produce plots of simulated wage profiles for an average woman considering alternatives for her fertility timing. These figures provide a useful visual depiction of the effect of motherhood on wage profiles, depicting both fixed costs and flatter profiles.

The previous regressions constrained wage growth to be the same for all women. The next set allows greater flexibility and heterogeneity in underlying wage profiles, by using wage growth as the dependent variable, and allowing for individual differences in levels and slopes of age-wage profiles. The unit of observation is still a person-age, but the outcome variable is now the change in log wages over the next three years.[18] I consider the following specification:

$$\Delta\text{Ln}(\text{Wage}_{i,t}) = \beta_0 + \beta_1\text{FirstBirth}_{it} + \beta_2\text{FirstBirth}_{it}\cdot(\text{A1B-20})_i + \beta_3\text{X}_i + \delta_t + \varepsilon_{it} \qquad (6)$$

where X_i contains the usual control variables, and δ_t are fixed effects for age to capture nonlinear wage growth. The dummy for FirstBirth is set to one in the year t=year of first birth. β_1 is a measure of the effect of motherhood on wage growth, and β_2 captures the change in this effect by age at first birth. A negative coefficient on β_1 indicates a reduction in earnings growth associated with motherhood, and a

[18] Using two-year and four-year intervals yielded similar but smaller coefficients.

positive coefficient on β_2 represents lower depreciation for older mothers. To account for changes in work hours surrounding birth, specifications are also estimated controlling for hours worked during the period.

The change equation form is similar to a regression on Ln(Wage$_{it}$) with individual fixed effects, as the base level of earnings is differenced out. This form is more flexible, though, since it allows wage growth to vary with individual factors, and for the impact of first birth to vary with its timing. In a related specification, individual fixed effects α_i are included to absorb personal idiosyncrasies in wage profile not accounted for by education, race, aptitude and birth year cohort. In addition, to allow for a lasting effect of motherhood on wage growth, I estimate the equation:

$$\Delta\text{Ln}(\text{Wage}_{i,t,t+3}) = \beta_0 + \beta_1\text{FirstBirth}_{it} + \beta_2\text{FirstBirth}_{it}\cdot(\text{A1B-20})_i + \beta_3\text{Mother}_{it} + \delta_t + \varepsilon_{it} \qquad (7)$$

which contains an additional term for motherhood, with and without individual fixed effects α_i. Controls for factors that affect wage growth improves on previous panel studies (e.g. Taniguchi, 1999), but the problem of endogenous fertility remains. I address this by estimating Equation 7 with the instrumental variables: $1(Age \geq \widehat{A1B})$, $1(Age+1 \geq \widehat{A1B} \geq Age-1)$, and $1(Age+1 \geq \widehat{A1B} \geq Age-1) \bullet (\widehat{A1B}-20)$, using bootstrapping to obtain standard errors.

5.3 Joint Estimation of Fertility Timing and Career Outcomes

Finally, I measure the converse effect of potential earnings on fertility timing. The first two results sections will establish the set of career financial incentives that women face in choosing optimal timing. The third section evaluates the responsiveness of timing choices to these incentives. The semi-log form of the wage equation (Equation 3) implies that the absolute financial benefit from motherhood delay $\beta_{A1B}\cdot(\beta_0+\beta_X X_i+\varepsilon_i)$ is increasing in earnings potential, even if the proportional benefit is the same for all women (β_{A1B}), producing a larger "price" effect for higher earners. The income effect from higher earnings could point in either direction, as wealthier women may choose to have children sooner, and the net effect is an empirical question. I use three-stage least squares to jointly estimate the following system of equations:

Timing of Motherhood Equation

$$\text{A1B}_i = \gamma_0 + \gamma_Y \text{Ln}(\text{Y}_i) + \gamma_M \text{Miscarriage}_i + \gamma_C \text{Contraception}_i + \gamma_L \text{YearsToBirth}_i + \gamma_X \text{X}_i + \varepsilon_{1i} \quad (8)$$

Earnings Equation

$$\text{Ln}(\text{Y}_i) = \beta_0 + \beta_{A1B} \text{A1B}_i + \beta_{AFQT} \text{AFQT}_i + \beta_X \text{X}_i + \varepsilon_{2i} \quad (9)$$

where Y_i represents either earnings or wage rates, X_i contains education, race and cohort, and the error terms ε_{1i} and ε_{2i} may be correlated. The exclusion restriction that identifies γ_Y is that AFQT matters for potential earnings but not for fertility timing; its validity is discussed in Section 7.3.

6. Data Description

This study uses data from the main NLSY79 public use files (1979-200).[19] The survey includes questions on a wide range of topics, including family background, education, beliefs, fertility and work histories, and annual earnings. It also features a panel structure, spanning twenty years, making it an appealing choice for labor market studies. For present purposes, this source is unusual in combining detailed labor market data with detailed pregnancy, childbirth and contraceptive use histories. The analysis excludes the military, minority and disadvantaged white over-samples;[20] the remaining respondents constitute a nationally representative sample of over 12,000 men and women, who were between the ages of 14 and 22 at their first interviews. After 1979, they were interviewed annually until 1994, and then biennially. Since the NLSY is a well-known survey, this section will focus on aspects particular to this paper: sample restrictions and variable construction.

[19] Confidential Geocode data are used for state level identification to test for differences in the tradeoff by family leave regime.
[20] The military and the economically disadvantaged sample were survey until 1983 and 1988, respectively.

To study the effects of motherhood timing, the sample is restricted to women with at least one child born during the survey period. The sample is further limited to those who had their first birth between the ages of 20 and 33, and between the years 1983 and 2000, excluding teen mothers (studied in Hotz et al., 1999) and childless women. Several other restrictions are motivated by data limitations: earnings observations are available for all birth year cohorts for the age range from 21 to 34, and contraceptive information starts in 1982. The number of potential observations is reduced to about 1,500, and the actual number is lower due to missing variables and sample attrition. Summary statistics are reported for the main variables in Table 1.

The dependant variables used in the empirical analysis are, in logarithmic form: net present value of career earnings (between ages 21 and 34), total career hours worked, average career wage rate, hourly wage rate for each age between 21 and 34, and change in hourly wage rates. For the period in which the survey is biennial, wage rates for missing years are imputed with linear interpolation.[21] Annual earnings over the period are aggregated into a net present value measure using inflation rates from the Bureau of Labor Statistics calculation of the Consumer Price Index. Results are not sensitive to using the Federal Reserve discount rates or Treasury bill interest rates (1-year risk-free rate) instead. For robustness, career earnings are also calculated over expanded windows: from age 21 to 35, 21 to 36, and on, until 21 to 42. The longer windows include smaller samples of women, but yield substantially similar results for career effects. However, one should remain cautious about extrapolating the results beyond the observed ages to predict lifetime effects, or beyond the sample to women becoming mothers after age 33. Hours worked each year are computed using actual hours reported in the work history files, and averaged over missing weeks within a year. Weeks of non-participation are included as zero hour weeks. The average wage rate is obtained by dividing the career earnings from 21 to 34 by the number of hours worked during that

[21] Although earnings and wage rates are only recorded for the year prior to each interview, the work history files cover all weeks from 1979 to 2000.

period.[22] The primary measure used for wage rates is the self-reported hourly wage rate at the primary job.

The basic controls included in most regressions are: birth year cohort, race, education level, and general aptitude. Date of birth and date of interview are reported in the survey. Race is summarized by indicator variables for Black and Hispanic, with the omitted category consisting primarily of White. Educational attainment is represented by dummy variables for High School Diploma and College Degree. The omitted category is high school dropouts, and College Degree also includes women with advanced and professional degrees. Armed Forces Qualification Test (AFQT) percentile score achieved in the 1980 test measures aptitude. Each woman is assigned to an occupation group based on predominant and most recent employment, among the categories: professional, technical and managerial; sales and service; clerical and kindred; and crafts, operatives, laborers, farmers and foremen. The family leave regime indicator is set to 1 for women who bore their first child in a state-year with a mandated protection rule. This applies to all women who bore their first child after the 1993 federal Family and Medical Leave Act, which provides up to twelve weeks of unpaid leave following the birth of a child (or adoption, or a medical condition affecting oneself or a family member), as long as the employer size and previous tenure conditions are met.[23] In addition, several states passed laws offering comparable protection prior to 1993.[24] When state laws are used for interactions, state fixed effects are included as controls, and identification is provided by state law changes during the sample period.

Survey questions about pregnancy outcomes provide data for the first instrumental variable, an indicator for miscarriage at first pregnancy. For women whose first pregnancy ended in abortion, the outcome of the first non-aborted pregnancy is used.[25] The indicator variable for contraceptive use at time

[22] This is equivalent to a weighted average, by hours worked at that rate, of all wage rates received during the period.
[23] See Ruhm (1997) for more background regarding the provisions, coverage, and consequences of the Family and Medical Leave Act.
[24] Waldfogel (1999) is the source for state maternity leave rules. States that adopted between 1972 and 1992 are, in chronological order of adoption: Massachusetts, Connecticut, Washington, California, Minnesota, Rhode Island, Maine, Oregon, Tennessee, Wisconsin, New Jersey, and Vermont.
[25] See Hotz et al. (1997) for a thorough discussion of how abortion affects the validity of miscarriage as an instrumental variable for fertility timing. Estimated bounds that incorporate potential contamination of the natural experi-

of pregnancy is taken from a combination of two questions asked after each live birth. The first asks if the woman used contraception prior to the pregnancy, and the second, if she ceased contraception prior to conception.[26] The indicator is intended to identify women who first became pregnant "accidentally" and in spite of precautionary efforts. The third instrumental variable, time to first conception, is constructed using biennial information on contraceptive use. Among the group of women who report using contraception at some point before their first birth, a woman's first conception attempt is defined to start at the first year that she reports sexual activity and no contraceptive use. The lag is defined as the number of years between the woman's first attempt and the birth of her first child. For women who never report contraception or whose first attempt is identified after first birth, the variable is set to zero. A companion dummy variable for "reported contraception prior to first birth" is included to remove the potential bias from contraceptive use. In addition, to account for maternal behaviors associated with an increased risk of miscarriage, control variables are added for substance use during first (or earliest reported) pregnancy, such as smoking cigarettes, drinking alcohol, and using marijuana or cocaine.

The incidence of miscarriage at first pregnancy in my sample is 14%, which is in line with Regan (2001), who cites medical estimates of about 15% of all recognized pregnancies ending in miscarriage. While a direct comparison of contraceptive use variables is not available, some general features of the NLSY match data from independent sources such as the Alan Guttmacher Institute and the National Surveys of Family Growth (NSFG). The most popular type of contraception for childless women is the birth control pill, although the pill has been somewhat replaced by condoms for women in more recent cohorts. The dummy variable for unintended pregnancy has a mean of about 0.30 in my sample. This may seem high when compared with failure rates of properly used contraceptive methods, but is lower than the

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ment from misreporting, and from latent abortion-types who miscarry, confirm the qualitative findings of their instrumental variables estimation. The problem of latent abortion-types in the miscarriage group is smaller in this paper, since the sample is limited to women who become mothers after age 20, and abortion rates are substantially lower (less than half as high) for later pregnancies

[26] The text of the first question reads: "Before you became pregnant the 1st time/between (date) and (date) did you ever use any methods to keep from getting pregnant?" The second reads: "Had you stopped all methods before you became pregnant?"

NSFG reported rates of unintended pregnancy (57% of pregnancies in 1987).[27] The NLSY79 also asks if pregnancies were "desired." The fraction of children who are "intended" is quite low. For example, consider responses to the 1982 question (R0769500), "Just before you became pregnant the first time, did you want to become pregnant when you did?" 11.6% of respondents said "Yes," 4.2% said "Didn't matter," 63.8% said "No, not at that time," and 20.5% said "No, not at all." These comparisons provide some corroboration for the accuracy of reported miscarriage and contraceptive use in the NLSY79, and support the claim that biological shocks play a role in human fertility.

7. Results

7.1 The Effects of Motherhood Delay on Earnings, Hours Worked, and Wage Rates

This section describes the measured effects of motherhood delay on women's career earnings, hours worked and wage rates, as well as terminal wage rates. Results are presented from estimating Equation 3 separately for each of the outcome measures with no discounting. Table 2 reports career outcomes: delayed motherhood leads to higher earnings and wage rates and more hours worked. OLS and IV estimates attribute an almost 10% increase in earnings per year of fertility deferment. Later mothers tend to have fewer children by the age of 34 than earlier mothers. To account for this potential difference in labor supply, number of children was added as a control, reducing the coefficient on A1B to about 6%, with a standard error of 0.025.[28]

If labor force participation is not affected by fertility timing, as in the fixed-interruption cases presented in the conceptual framework, total earnings and average wages should be equivalent outcome measures. However, the IV results for Ln(Hours) indicate that later mothers tend to work more hours during the period. This can be an artifact of the early cutoff (at age 34, where more of the late mothers have not completed their interruptions), or the result of differences in the labor market supply and demand

[27] The difference in rates of accidental pregnancy and accidental live birth can be explained by abortion.

[28] Also, to account for endogeneity in number of children, the variable was also instrumented for with the main IVs, leading to a non-significant negative effect for number of children, and a coefficient estimate of 0.061 (standard error=0.037) for A1B.

for work by new mothers of different ages and levels of experience. One possibility is that older first-time mothers are more likely to secure childcare arrangements for their infants and return more quickly to full-time employment.[29] Because of the systematic difference in hours worked, the results for earnings confound several factors. Early mothers earn less, but also consume more leisure and home time. For this reason, it may be easier to interpret results for the average wage rate. The estimates show a 3% increase in wages per year of deferment with a standard error of 0.010.

Additional regressions including controls for marital history and birth year dummies were not reported in the table, but showed similar patterns. Over 90% of the women in the sample were married prior to their first birth and 56% were continuously married. After controlling for these two factors, the IV point estimate of A1B for earnings increases to 12% and average wage rate to 3.6%, both significant at 1%. An artifact of the data selection rule is that women from earlier birth cohorts are over-represented among the older mothers. To ensure that the A1B results are not driven by non-linearity in the effect of birth year, regressions were run using birth year dummies, again leading to substantially similar results: A1B coefficients of 10.2% for earnings and 3.3% for wages, significant at 1%.

Table 3 reports estimation of the first stage regression (Equation 2) and of an equation with separate variables for “accidental” pregnancy by contraception type. Coefficients on the instrumental variables listed in the first column are reasonable in sign and magnitude: Miscarriage defers first birth by about 6 months, and unintended pregnancy brings first birth 8.5 months earlier. A year of conception lag is associated with 9 months of delay, yielding an average delay of 1 year and 2 months at the sample mean of 1.5. Typically, only one of the IVs applies to a given woman, so the instruments are causing about 6 months to a year of fertility shifts. These are small, but not trivial effects that provide identification in the second stage, where the A1B coefficient should be interpreted as the average effect of

[29] In separate Probit and IV-Probit results, age at first birth is found to be a positive predictor of working full time in the year, 2 years and 3 years following first birth.

a year of motherhood delay due to biological fertility shocks.[30] The second column shows that failed contraception predicts motherhood delay best for women using contraceptive foam, or taking the birth control pill, the most popular form of contraception for young childless women in the NLSY79. It is worth noting that the instruments are statistically significant predictors of A1B and the F-statistic on the joint significance test is F(3,1284)=160. Results from tests of the over-identifying restrictions are reported last rows of Table 2. Exogeneity is not rejected at the 10% significance level, although the statistic is borderline in Column 7 where log-hours is the dependent variable.[31]

Regressions that allow for heterogeneous effects of A1B are reported in Table 4. Column 2 reports IV results by occupation type. Motherhood delay has the largest, and statistically significant, benefit on the average wage rates of women in professional or managerial occupations (coefficient of 0.047), followed by those in clerical occupations (coefficient of 0.022). Sales and service workers have negligible returns (coefficient of 0.008), and women in crafts and manual labor have marginally significant negative returns (coefficient of -0.098, with a standard error of 0.055). The A1B·(Prof/manager) is significantly different from A1B·(Craft/labor) and A1B·(Sales/service), although not significantly different from A1B·Clerical. By contrast, the OLS regression yields significant A1B coefficients for women in professional/managerial, clerical, and sales/service occupations, with the first term actually lower than the IV estimate. Compared to the results in Table 2, where OLS and IV results were nearly identical, this table suggests a more complex story for selection by A1B: there appears to be negative selection within professional and managerial occupations and positive selection for the others. Column 4 contains IV estimates by education level, where college educated women exhibit substantial returns to motherhood delay (of about 5.3% wage benefit per year of delay) that are significantly larger than the returns for women with less formal education (at the 5% significance level). Here, selection is negative for the college graduates and positive for others. The results by education and by occupation

[30] Since shocks occur during the entire period, the estimate captures the average effect of a delay across women of different ages. Regressions were run using polynomial terms in A1B to consider non-linearity in the effect of delay by A1B, but coefficients were imprecisely estimated.
[31] In a related exercise, the regressions were repeated omitting each of the instruments in turn. Coefficient estimates were qualitatively identical and statistically indistinguishable

confirm the prior belief that motherhood delay is most important for women working in careers characterized by ongoing human capital accumulation and costs for interruptions.

The IV interaction terms for government employment (in Column 6) and spouse earnings (unreported) are both negative, but are not statistically different from zero. Family leave regime does not appear to influences wage rates or returns to delay either (not reported). The OLS coefficient on A1B·FamilyLeave is -0.003 with a standard error of 0.009, and the IV estimate is –0.010 with a standard error of 0.020.[32] The coefficients on FamilyLeave range from –1.11 (standard error of 1.20) to 0.44 (standard error of 0.53). The direct effects of family leave and interactions with A1B are imprecisely estimated and qualitatively inconsistent, providing no supporting evidence for the role of mandated leave in mitigating the conflict between career and early motherhood.

As discussed in Section 3 (Case 4, Figure 7), in the special case where the only effect of motherhood is a fixed cost of human capital depreciation, fertility timing affects career earnings and average wages, but not terminal wage rates. Table 5 contains results from estimating Equation 4 with wages at age 34, or "terminal" wage rates, as the dependent variable. The OLS coefficient on A1B implies that a year of fertility deferment is associated with a 4.1% increase in wages, statistically significant at the 1% level. The baseline IV estimate in Column 2 is lower, around 2.6% and is less precisely estimated than OLS. Columns 3 and 4 show how the coefficient responds to the inclusion of experience hours as a control, first as an exogenous regressor, and then as an endogenous variable, using the core instrumental variables. The first change increased the estimate of the A1B coefficient, while the second reduced it, and left it statistically insignificant.

The final column shows how a Heckman two-step correction for sample selectivity changes the results. In this case, the probability of working at age 34, and of generating a wage rate observation, was predicted in a Probit regression using all of the controls from the main regression, and the additional participation instruments of marital status and spouse earnings. Some unreported variations were

[32] In another specification, state laws were ignored, and family leave was presented by an indicator for first birth occurring after the Family and Medical Leave Act (FMLA). The coefficient on A1B*(FMLA) is 0.018 in OLS and 0.037 in IV.

attempted, such as treating A1B as endogenous in the first step, and including higher order polynomial terms of the Inverse Mills Ratio in the second step. Across specifications, the A1B coefficient was consistently positive and around 3.5%. The positive effect of delay on terminal wage rates suggests that depreciation costs are lower for older mothers or that women experience a reduction in wage growth following motherhood. In addition, when a similar IV specification was used to test for the impact of A1B on wages *prior* to motherhood (wages at ages 22 and 25 for women who become mothers after ages 22 and 25, respectively), the coefficients were negative (-0.01 and -0.026) and statistically insignificant (standard errors 0.011 and 0.025). While certainly not conclusive, this does suggest that the instruments are not operating through a correlation with fixed omitted factors.

7.2 The Effects of Motherhood on Wages and Wage Growth

The next set of empirical exercises directly estimates the wage effects of motherhood by further exploiting the panel nature of the data. Table 6 contains results from the three primary specifications for log-wages: linear, quadratic (Equation 4), and quadratic wage growth with linear costs of motherhood that vary with first birth timing (Equation 5). As explained in Section 5.2, standard errors for the IV coefficients, in Columns 4 to 6, are estimated by bootstrapping with 1,000 replications. Mother and Mother·YearsSinceFirstBirth have substantial negative coefficients, indicating that mothers suffer both fixed costs and lower wage growth. In Column 6, the fixed costs of motherhood are held constant and changes in slope for mothers are allowed to vary by timing of first birth. The coefficient for the women with $20 \leq A1B < 25$ is more negative than for the $25 \leq A1B < 30$ group, and the coefficient for the oldest group of mothers is most negative but imprecisely estimated. Figures 9 and 10 show plots of potential age-wage profiles, based on the IV coefficients in Columns 3 and 5. The solid line represents the wage profile for a childless woman, while the dashed and dotted lines show the expected profile for that same woman, with A1B=25 and A1B=30. The key elements apparent in the figures are that women: (1) return at lower wages and (2) experience lower growth in wages following first birth. Together, these factors produce the career and terminal wage benefits from motherhood delay estimated in Section 7.1.

The next approach using panel data takes increase in log-wages over a 3-year period as the dependent variable and improves on two weaknesses of the previous approach. First, I relax the assumption that all women experience the same wage growth, and second, I consider actual experience as a driver of growth. The first four columns of Table 7 report results from specifications in which education, race, AFQT and birth cohort separately affect wage growth, while the last four columns are from specifications allowing wage growth to vary by individual. Columns 2 and 6 report results that control for actual experience. All regressions include age fixed effects to flexibly capture nonlinear wage growth.

The first specification corresponds to Equation 6, which constrains mothers to be on the same age-wage profile as non-mothers, but allows for local motherhood costs in the period immediately following first birth, measured by the FirstBirth coefficient. The estimates are easiest to interpret under the model with no permanent slope change from motherhood in Cases 3 and 4 of Section 3, illustrated in Figures 6 and 7. In Columns 1 and 5, FirstBirth captures fixed motherhood penalties as well as potential costs of foregone human capital accumulation during a labor market absence. When I also control for hours worked, the coefficient on FirstBirth is somewhat smaller and still significantly different from zero, suggesting that fixed costs dominate foregone experience in causing local motherhood costs. The negative coefficient on FirstBirth loosely corresponds to the negative coefficient on Mother in the previous table; a temporary reduction in wage growth causes a permanent reduction in wage levels. In Table 7, the FirstBirth·(A1B-20) term allows FirstBirth costs to vary with motherhood timing. The positive coefficients (around 1% in OLS and 8% in IV) indicate that costs are lower for older mothers.

Next, a Mother indicator is added to accommodate a permanent change in the slope of the age-wage profile, following Equation 7 (and combining features from Cases 4 and 5 of Section 3). The coefficient is only significantly different from zero in Column 3, but the negative sign across columns suggests a flattening of wage profiles for mothers. Adding the Mother variable does not significantly change the FirstBirth or FirstBirth·(A1B-20) coefficients (comparing Column 1 to Column 3, or 5 to 7). Finally, I address the potential endogeneity of motherhood timing by estimating Equation 7 using IV with and without individual fixed effects (Columns 4 and 8). Although the FirstBirth and FirstBirth·(A1B-20) coefficients are much larger and less precise than in OLS, the IV results do reproduce the qualitative features and approximate relative magnitudes of the two coefficients, providing weak corroboration. The lack of precision may not be surprising considering the unusual dependent variable and that the variation in the core instrumental variables is entirely cross-sectional.

7.3 Joint Estimation of Fertility Timing and Career Outcomes

The previous two sections demonstrated the presence of labor market incentives for fertility delay; this section measures the fertility timing response. I will first present evidence supporting the validity of AFQT as an instrument for potential earnings, and I will then discuss the results of joint estimation of fertility timing and career outcomes.

AFQT percentile is used as an instrumental variable for earnings in the fertility timing equation. The necessary exclusion restriction is that AFQT does not independently or directly belong in the fertility timing equation, that is, conditional on race, birth year cohort, and education level, that a woman's intelligence is uncorrelated with her preferences for motherhood timing. While this assumption apparently contradicts the empirical fact that AFQT is a strong predictor of fertility timing, the observed correlation can be due to the indirect effect of AFQT operating through potential wages. To determine the relative importance of the direct and indirect effects, I subdivide the sample into three types of women, by their labor force participation status between ages 20 and 34, and I estimate the following equation:

$$\begin{aligned} \text{A1B}_i = {} & \beta_F \text{AFQT}_i \cdot \text{FullTimer}_i + \beta_P \text{AFQT}_i \cdot \text{PartTimer}_i + \beta_H \text{AFQT}_i \cdot \text{HomeTimer}_i \\ & + \beta_X \text{X}_i + \varepsilon_i \end{aligned} \qquad (10)$$

where the effect of AFQT is entered interacted with worker type: largely full-time, part-time, and home-time, and X_i includes indicators for FullTimer, PartTimer and HomeTimer. The estimated coefficients are

$\beta_F = 0.027$ (standard error of 0.007), $\beta_P = 0.007$ (standard error of 0.005), and $\beta_H = -0.003$ (standard error of 0.006). The trend of decreasing coefficients from full to home groups and the negative coefficient on the HomeTimer group (about a third of the sample) indicate that AFQT does not directly affect fertility timing: as the career channel is removed, the association between AFQT and A1B disappears. A potential weakness of this exercise is that women select into their employment groups, and that tastes for family might be correlated with this selection. Controlling for number of children does not change the qualitative results, suggesting that the sorting is being driven by variation in outside wealth and tastes for leisure, in which case, the results of this exercise support the claim that AFQT does not directly affect fertility timing.

I employ a simultaneous equations framework for estimating the joint effects of timing on earnings and earnings on timing. Results from the three-stage least squares estimation of Equations 8 and 9 are reported in Table 8. The effect of a year of fertility delay on earnings is about 8%, and the earnings coefficient on A1B is 1.546 with a standard error of 0.397. This converts to a 67% increase in earnings causing a single year of fertility deferment. In terms of the observed wage distribution, an exogenous shift from the bottom quartile (log-wage = 11.94) to the top (log-wage = 12.98), will lead to (12.98 - 11.94)*1.55 = 1.61 years of deferment.[33] The effect is statistically significant, and magnitudes are non-trivial. The simplest interpretation is that women do consider lifetime earnings in timing their fertility; higher expected earnings lead to motherhood delay. An alternative story that is also consistent with the fact that AFQT only predicts A1B for working women is that higher ability women work in jobs that provide more intrinsic satisfaction, and are thus more willing to delay motherhood for their careers. Differences in tastes for children that are correlated with AFQT but unrelated to job characteristics are a far less plausible explanation.

[33] A similar shift from the bottom to top decile will add 3.667 years of deferment.

8. Conclusion

This research shows that fertility timing affects career average female earnings, hours worked and wages, as well as post-motherhood wage rates. Fertility delay can reduce the "family gap" in pay. The direct panel evidence indicates a likely channel for the effect: a "mommy track." Women experience reduced earnings around the time of their first birth and a flattening of the age-wage profile following motherhood. Whether this is caused by mothers choosing to reduce their labor supply and human capital investment, or by employers offering differential treatment with diminished training and advancement opportunities to mothers, cannot be determined empirically from the current evidence. In fact, the two are likely to be interconnected. However, the importance of wage growth and, by implication, of human capital investment, is echoed in the heterogeneity of the fertility timing effect: fertility delay improves career outcomes only for women with college degrees and those in professional or managerial occupations. As expected, women in careers with greater wage growth are the ones who gain financially from delaying motherhood.

Turning to the converse question, this paper shows that higher expected earnings lead to later motherhood. The career benefits of delay are greater for women with higher earnings, while the costs of delay – medical, emotional, and other – do not seem to increase with earnings enough to fully offset the benefits. Taken together, the results show that motherhood timing is an important dimension of the career-family conflict for many women, and that long-term career factors can matter on the fertility planning horizon.

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Table 1: Summary Statistics

Variable	Observations	Mean	Std. Dev.
Age at first birth	1,595	26.678	3.586
Age at first pregnancy	1,556	24.514	4.286
First pregnancy ended in miscarriage	1,573	0.137	
Contraception at first pregnancy	1,571	0.295	
Lag before first birth (years)	1,595	1.535	2.664
Positive lag before first birth	1,595	0.384	
Never divorced or separated	1,595	0.559	
Married before first birth	1,415	0.912	
NPV earnings 21-34 (in 1999$)	1,212	309,331	203,489
Hourly wage rate at 21 (in 1999$)	1,356	4.956	3.730
Hourly wage rate at 34 (in 1999$)	1,066	9.722	17.958
Average weekly hours (age 21-34)	1,579	29.095	13.041
Number of children	1,595	2.111	0.924
Number of children expected (in 1979)	1,578	2.493	1.451
Black	1,595	0.176	
Hispanic	1,595	0.224	
No high school diploma	1,567	0.057	
High school diploma	1,567	0.522	
College degree or higher	1,567	0.421	
Alcohol during pregnancy	1,582	0.516	
Smoking during pregnancy	1,582	0.243	
Cocaine during pregnancy	1,472	0.012	
Marijuana during pregnancy	1,472	0.030	

Sample: NLSY 1979 women who report at least one child by 2000, 1st birth between ages of 21 and 34, during the period from 1983-2000.

Table 2: Effects of Fertility Timing on Career Earnings, Wage Rates, and Hours Worked

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Method:	OLS	IV	IV	OLS	IV	OLS	IV
Dependent Variable	*Ln(Earnings)*			*Ln(WageRate)*		*Ln(Hours)*	
A1B	0.097**	0.099**	0.059*	0.032**	0.033**	0.061**	0.047**
	[0.010]	[0.021]	[0.025]	[0.005]	[0.011]	[0.006]	[0.013]
Birth Year Cohort	-0.102**	-0.106**	-0.101**	-0.060**	-0.062**	-0.036**	-0.033**
	[0.016]	[0.017]	[0.017]	[0.008]	[0.009]	[0.010]	[0.011]
HS diploma	1.417**	1.398**	1.435**	0.284**	0.268**	0.700**	0.710**
	[0.160]	[0.160]	[0.157]	[0.087]	[0.086]	[0.091]	[0.091]
College or higher	1.567**	1.549**	1.597**	0.439**	0.421**	0.716**	0.751**
	[0.173]	[0.178]	[0.176]	[0.093]	[0.095]	[0.100]	[0.104]
Black	0.198*	0.181*	0.151+	0.104*	0.101*	0.091	0.073
	[0.091]	[0.091]	[0.090]	[0.048]	[0.048]	[0.056]	[0.057]
Hispanic	0.338**	0.326**	0.327**	0.17**	0.165**	0.126*	0.119*
	[0.094]	[0.095]	[0.093]	[0.049]	[0.050]	[0.056]	[0.057]
AFQT percentile	0.009**	0.008**	0.009**	0.006**	0.006**	0.004**	0.003**
	[0.002]	[0.002]	[0.002]	[0.001]	[0.001]	[0.001]	[0.001]
Report contraception		-0.146*	-0.11		-0.041		-0.076+
		[0.068]	[0.068]		[0.036]		[0.043]
Smoking		-0.146+	-0.161*		-0.079+		-0.084+
		[0.077]	[0.076]		[0.041]		[0.049]
Alcohol		0.129+	0.108+		0.083*		0.072+
		[0.066]	[0.065]		[0.034]		[0.042]
Marijuana		-0.08	-0.147		-0.157		0.109
		[0.209]	[0.206]		[0.109]		[0.131]
Cocaine		-0.679+	-0.602+		-0.243		-0.560**
		[0.371]	[0.365]		[0.194]		[0.200]
# Children by age 34			-0.26**				
			[0.052]				
Observations	1,019	1,019	1,019	1,016	1,016	1,282	1,282
R-Squared	0.28			0.24		0.17	
Over-identification test of all instrumental variables							
Sargan statistic		2.017	0.523		1.743		4.584
P-value		0.365	0.770		0.418		0.101

Note: Dependent variables are calculated over the age range from 21 to 34, with no discounting. Regression 3 includes number of children by age 34 as an independent variable. Sample includes all women with full experience profiles, who had their first child between the ages of 21 and 33, and the years 1983-2000.

Standard errors in brackets
+ significant at 10%; * significant at 5%; ** significant at 1%

Table 3: First Stage Regressions on Age at First Birth

Dependent Variable: Age at first birth

	Basic	Contraception Type
First pregnancy ended in stillbirth/miscarriage	0.514* [0.202]	0.808** [0.234]
Years to conception, first child	0.750** [0.036]	0.659** [0.042]
First pregnancy while using contraception	-0.714** [0.156]	
Pregnant using: birth control pill		-0.707** [0.222]
Pregnant using: condom		-0.104 [0.396]
Pregnant using: foam		-1.824* [0.893]
Pregnant using: jelly/cream		0.248 [1.559]
Pregnant using: suppository/insert		-1.367 [1.092]
Pregnant using: diaphragm		-0.591 [0.639]
Pregnant using: douching		-0.526 [1.885]
Pregnant using: IUD/coil/loop		-0.492 [1.260]
Pregnant using: natural family planning		0.655 [1.159]
Pregnant using: rhythm (by calendar)		-0.072 [0.731]
Pregnant using: withdrawal		-0.062 [0.822]
Pregnant using: abstinence, cervical cap, Norplant, other		0.249 [0.630]
Constant	16.800** [0.632]	10.559** [0.886]
Test of joint significance of instruments	F(3, 1284)= 159.16	F(16, 1055)= 17.97
Observations R-squared	1,299 0.44	1,092 0.40

Note: OLS regressions, coefficients not reported for birth year, race, education level, AFQT, contraception use prior to motherhood, and substance use indicators. The second regression breaks down the third IV by contraception type. Sample is all women with full experience profiles, who reported a first birth between the ages of 21 and 33, and the years 1983-2000.

Standard errors in brackets. * significant at 5%; ** significant at 1%.

Table 4: Heterogeneous Effects of Fertility Timing on Career Wage Rates

Dependent variable: Ln(Average wage rate from ages 21-34)=Ln(Earnings 21-34)-Ln(Hours 21-34)

Regression Method	(1) OLS	(2) IV	(3) OLS	(4) IV	(5) OLS	(6) IV
A1B·(Prof/manager)	0.027	0.047				
	[0.008]**	[0.019]*				
A1B·Clerical	0.022	0.022				
	[0.007]**	[0.016]				
A1B·(Craft/labor)	0.016	-0.098				
	[0.018]	[0.055]+				
A1B·(Sales/service)	0.044	0.008				
	[0.009]**	[0.021]				
A1B			0.032**	0.002	0.033**	0.027*
			[0.006]	[0.014]	[0.005]	[0.011]
A1B·(College grad)			-0.001	0.051*		
			[0.009]	[0.023]		
A1B·Government					-0.018	-0.012
					[0.016]	[0.050]
A1B·FamilyLeave						
Observations	1,105	1,016	1,105	1,016	1,105	1,016
R-squared	0.27		0.25		0.25	

Note: Controls included in the regressions, with unreported coefficients, are: constant term, education, race, AFQT, birth year cohort, as well as substance use variables, and dummy variables for occupation type and employment sector, as applicable.

Standard errors in brackets.
+ significant at 10%; * significant at 5%; ** significant at 1%

Table 5: Effects of Fertility Timing on Terminal Wage Rates

Dependent variable: Ln(WageRate$_{i,34}$)

	(1)	(2)	(3)	(4)	(5)
Method:	OLS	IV	IV	IV	IV
A1B	0.037**	0.026+	0.035+	0.018	0.038+
	[0.006]	[0.013]	[0.020]	[0.013]	[0.020]
HS Diploma	-0.009	-0.007	0.067	-0.082	0.119
	[0.103]	[0.103]	[0.160]	[0.100]	[0.218]
College or Higher	0.228*	0.244*	0.300*	0.180+	0.354
	[0.110]	[0.113]	[0.153]	[0.108]	[0.226]
Black	0.091+	0.081	0.066	0.098+	0.114
	[0.054]	[0.054]	[0.064]	[0.053]	[0.092]
Hispanic	0.150**	0.128*	0.141*	0.108*	0.178*
	[0.054]	[0.055]	[0.065]	[0.053]	[0.079]
AFQT Percentile	0.007**	0.006**	0.007**	0.005**	0.006**
	[0.001]	[0.001]	[0.002]	[0.001]	[0.001]
Birth Year Cohort	-0.074**	0.045	0.051	0.031	-0.064**
	[0.009]	[0.041]	[0.047]	[0.040]	[0.015]
Smoking		-0.09+	-0.08	-0.086+	-0.13+
		[0.049]	[0.054]	[0.047]	[0.075]
Alcohol		0.067	0.102	0.046	0.043
		[0.042]	[0.062]	[0.041]	[0.059]
Marijuana		-0.136	-0.166	-0.115	0.021
		[0.123]	[0.140]	[0.118]	[0.192]
Cocaine		-0.009	-0.108	0.077	0.091
		[0.212]	[0.271]	[0.205]	[0.405]
Report Contraception		-0.046	-0.068	-0.039	-0.106+
		[0.044]	[0.053]	[0.043]	[0.063]
Inverse Mills Ratio					0.701**
					[0.201]
Experience Hours			-0.001	0.001**	
			[0.002]	[0.000]	
Constant	6.433**	-221.60**	-273.17*	-165.38*	5.906**
	[0.205]	[77.310]	[117.967]	[75.380]	[0.486]
Observations	1,187	1,187	1,170	1,170	674

Note: In regression 3, actual experience hours are included as an additional endogeneous variable (same IVs), while in 4 it is as treated as exogenous. Regression 5 uses a Heckman two-step procedure to correct for sample selectivity. Instruments for labor force participation are current spouse earnings and marital status.

Standard errors in brackets.
+ significant at 10%; * significant at 5%; ** significant at 1%

Table 6: Effects of Motherhood on Wage Profiles

Dependent Variable: Ln(Wage rate$_{it}$)

	(1)	(2)	(3)	(4)	(5)	(6)
Method	OLS	OLS	OLS	IV	IV	IV
Mother	-0.095**	-0.078**	-0.095**	-0.206**	-0.222+	-0.246**
	[0.012]	[0.014]	[0.013]	[0.060]	[0.126]	[0.063]
Mother·YearsSinceFirstBirth	-0.045**	-0.047**		-0.080**	-0.072**	
	[0.002]	[0.005]		[0.004]	[0.025]	
Mother·YearsSinceFirstBirth^2		0.001**			0.000	
		[0.000]			[0.003]	
Mother·YearsSinceFirstBirth·1(20<A1B≤25)			-0.033**			-0.079**
			[0.003]			[0.009]
Mother·YearsSinceFirstBirth·1(25<A1B≤30)			-0.039**			-0.047+
			[0.005]			[0.028]
Mother·YearsSinceFirstBirth·1(30<A1B≤35)			-0.004			-0.212
			[0.018]			[0.129]
Age	0.065**	0.177**	0.171**	0.091**	0.128**	0.131*
	[0.002]	[0.013]	[0.015]	[0.006]	[0.017]	[0.058]
Age^2		-0.002**	-0.002**		-0.001*	-0.001
		[0.000]	[0.000]		[0.000]	[0.001]
Observations	16,542	16,542	16,542	14,527	14,527	14,491

Note: Regressions include individual fixed effects. Samples include women with their first child born between the ages of 21 and 33 and years 1983-2000. Unit of observation is the individual-year from age 20-34. OLS Robust standard errors in brackets. IV standard errors calculated by bootstrapping with 1,000 replications.

+ significant at 10%; * significant at 5%; ** significant at 1%

Table 7: Effects of Motherhood on Wage Growth

Dependent variable: Change in Ln(WageRate) over 3 year period = $Ln(WageRate_{i,t+3})$-$Ln(WageRate_{i,t})$

Regression	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Method	OLS	OLS	OLS	IV	OLS	OLS	OLS	IV
Fixed Effects?	N	N	N	N	Y	Y	Y	Y
FirstBirth	-0.14**	-0.112**	-0.121**	-1.206	-0.107**	-0.090*	-0.106**	-1.262
	[0.036]	[0.036]	[0.037]	[1.334]	[0.037]	[0.037]	[0.037]	[12.42]
(A1B-20)·FirstBirth	0.013**	0.011*	0.011*	0.07	0.009+	0.008	0.009+	0.086
	[0.005]	[0.005]	[0.005]	[0.149]	[0.005]	[0.005]	[0.005]	[1.767]
Mother			-0.027*	-0.036			-0.007	-0.123
			[0.012]	[0.054]			[0.015]	[1.405]
HS diploma	-0.015	-0.032	-0.017	-0.021				
	[0.020]	[0.020]	[0.020]	[0.024]				
College degree	0.049*	0.028	0.045*	0.036				
	[0.022]	[0.022]	[0.022]	[0.036]				
AFQT percentile	0.001**	0.001**	0.001**	0.001**				
	[0.000]	[0.000]	[0.000]	[0.0002]				
Black	0.039**	0.034**	0.040**	0.039**				
	[0.012]	[0.012]	[0.012]	[0.015]				
Hispanic	0.013	0.008	0.013	0.016				
	[0.012]	[0.012]	[0.012]	[0.015]				
Birth Year Cohort	-0.026**	-0.025**	-0.027**	-0.028**				
	[0.002]	[0.002]	[0.002]	[0.005]				
Hours worked		0.001**				0.001**		
		[0.000]				[0.000]		
Constant	0.622**	0.559**	0.726**	0.583**	0.217**	0.147**	0.303**	0.25**
	[0.044]	[0.044]	[0.049]	[0.130]	[0.017]	[0.021]	[0.025]	[0.025]
Observations	19,033	19,033	19,033	17,171	19,033	19,033	19,033	17,171
R-squared	0.03	0.04	0.03		0.02	0.02	0.02	

Note: All regressions include a full set of age fixed effects. FirstBirth is a dummy variable set to one in the year of first birth. Sample includes women who reported a first birth between the ages of 21 and 33, and the years 1983-2000. OLS robust standard errors in brackets. IV standard errors calculated by bootstrapping with 1,000 replications.

+ significant at 10%; * significant at 5%; ** significant at 1%

Table 8: Joint Estimation of Fertility Timing and Career Outcomes

Dependent Variable	*Ln(Earnings)*	*A1B*	*Ln(Wage rate)*	*A1B*
A1B	0.084**		0.028**	
	[0.020]		[0.010]	
Ln(Earnings)		1.546**		
		[0.397]		
Ln(Wage rate)				2.359**
				[0.594]
Miscarriage/stillbirth		0.553**		0.588**
		[0.205]		[0.219]
Time to 1st conception		0.663**		0.712**
		[0.048]		[0.044]
Contraception at 1st preg.		-0.423**		-0.444**
		[0.157]		[0.161]
AFQT percentile	0.009**		0.006**	
	[0.001]		[0.001]	
HS diploma	1.32**	-1.389+	0.255**	0.016
	[0.149]	[0.733]	[0.079]	[0.474]
College or higher	1.492**	-0.672	0.408**	0.715
	[0.165]	[0.905]	[0.087]	[0.602]
Black	0.145+	-0.646**	0.040	-0.541*
	[0.084]	[0.210]	[0.044]	[0.221]
Hispanic	0.258**	-0.643**	0.13**	-0.536*
	[0.088]	[0.230]	[0.046]	[0.231]
Birth Year Cohort	-0.068**	0.575**	-0.027**	0.576**
	[0.016]	[0.039]	[0.008]	[0.039]
Reported contraception	-0.108+	-0.428*	-0.018	-0.598**
	[0.065]	[0.209]	[0.034]	[0.201]
Constant	8.681**	-0.543	0.860**	11.727**
	[0.443]	[4.038]	[0.232]	[1.111]
Observations	1,121	1,121	1,106	1,106

Note: results are from three-stage least squares estimation of earnings and timing (in regressions 1 and 2) and average wages and timing (in 3 and 4). Samples include women with their first child born between the ages of 21 and 33 and years 1983-2000.

Standard errors in brackets
+ significant at 10%; * significant at 5%; ** significant at 1%

Figure 1: Earnings Profiles by Age at First Birth

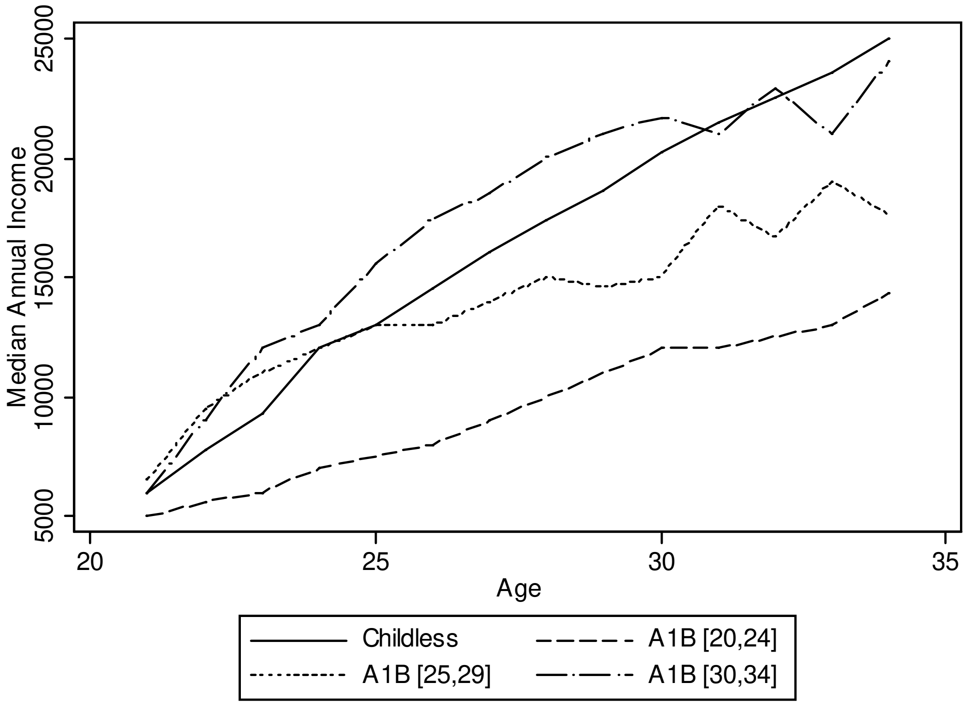
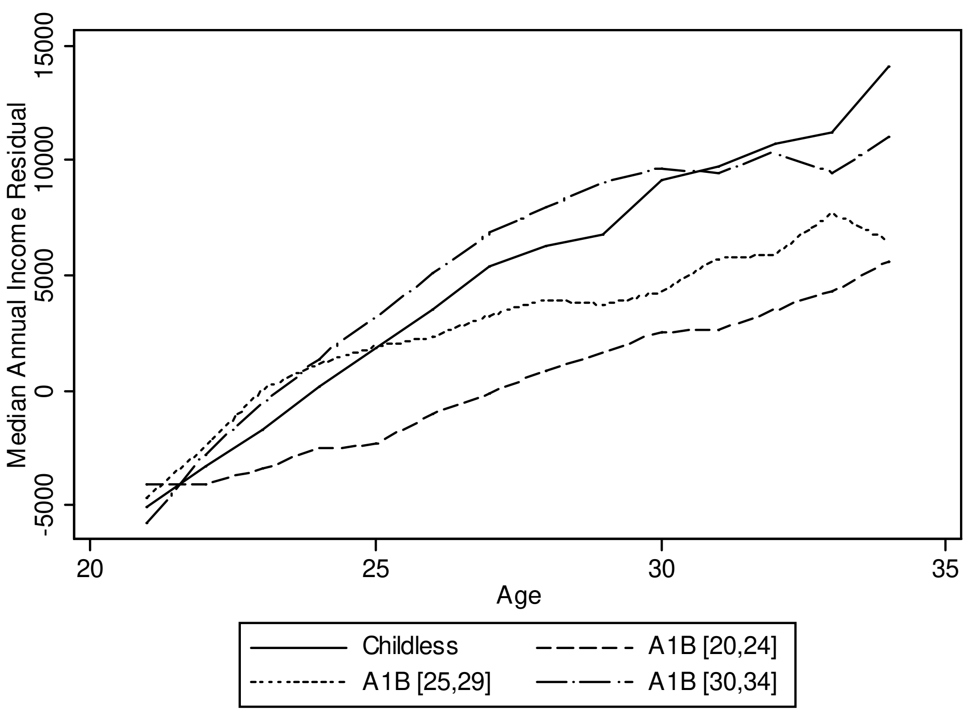


Figure 2: Earnings Residual Profiles by Age at First Birth



Note: Residuals are the difference between actual and predicted log-earnings from a linear regression with education, race, ability and birth year cohort as controls.

Figure 3: Hours Worked Profiles by Age at First Birth

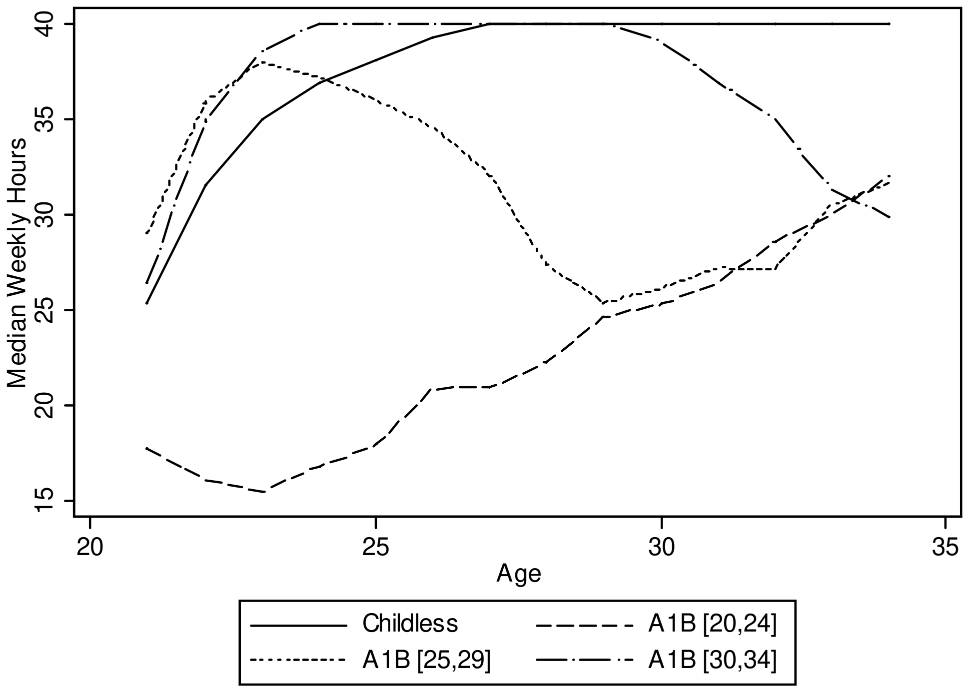


Figure 4: Wage Rate Profiles by Age at First Birth

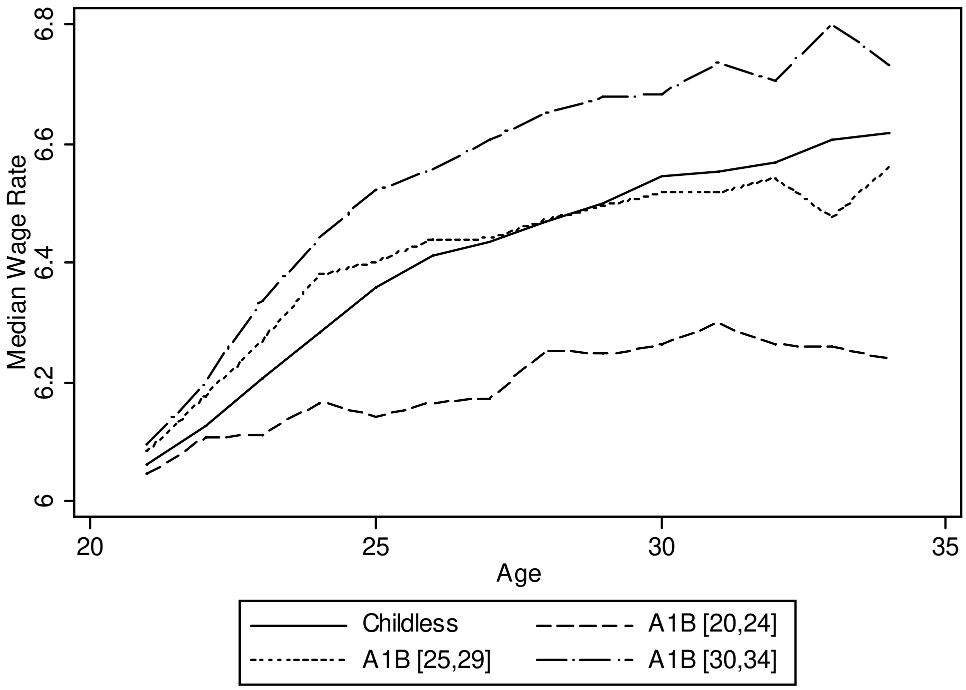
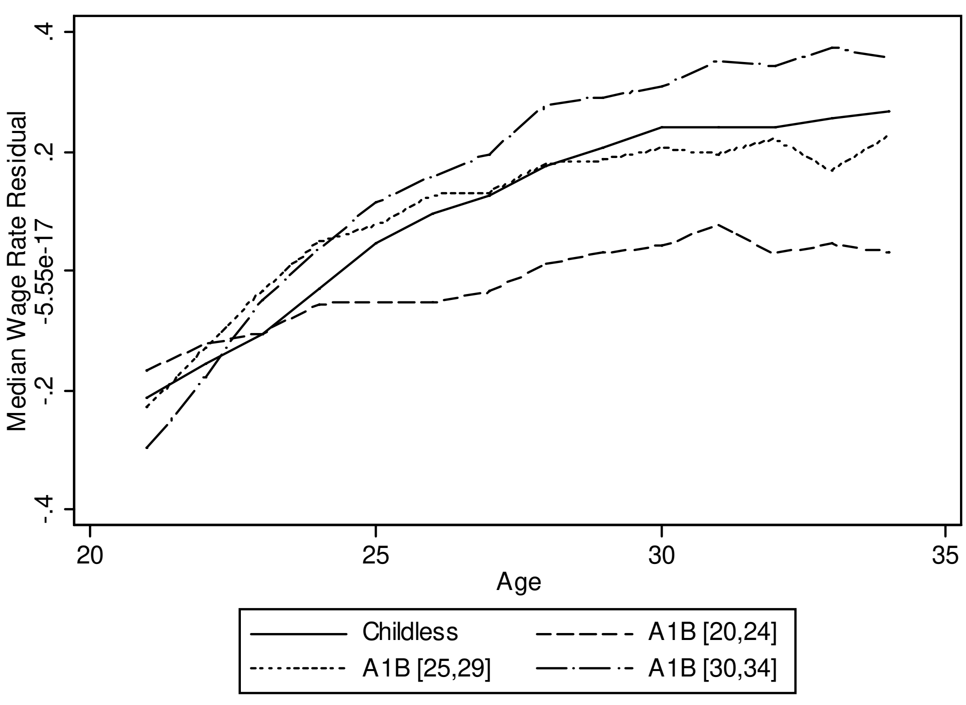
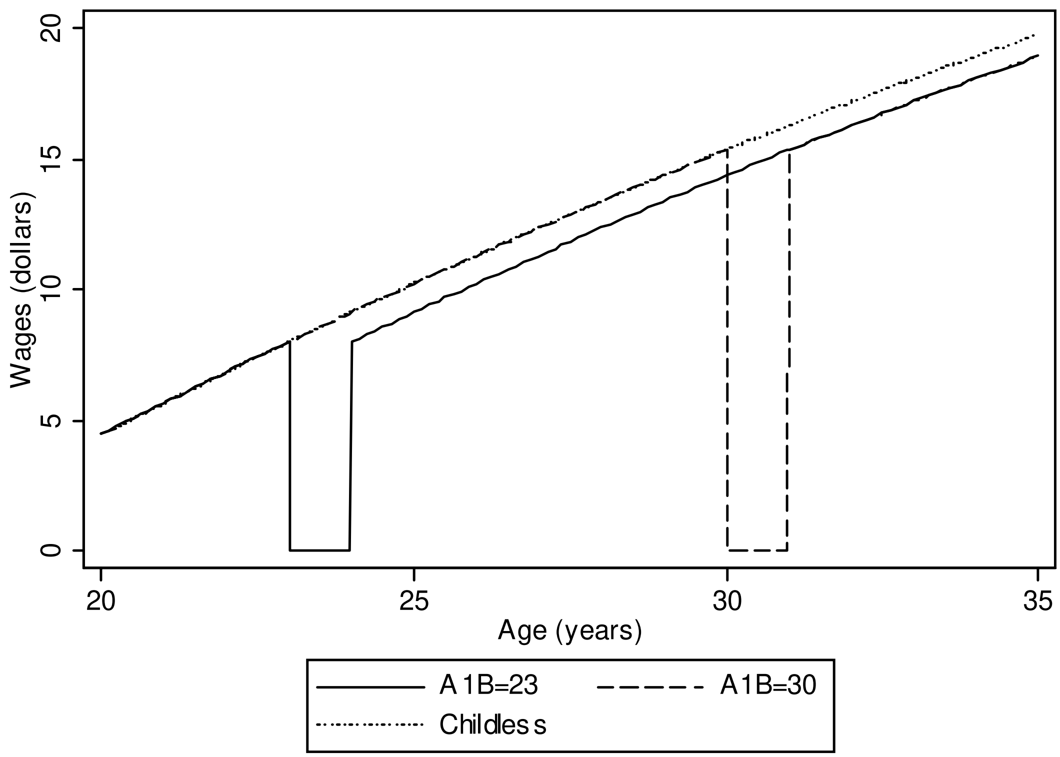


Figure 5: Wage Rate Residual Profiles by Age at First Birth



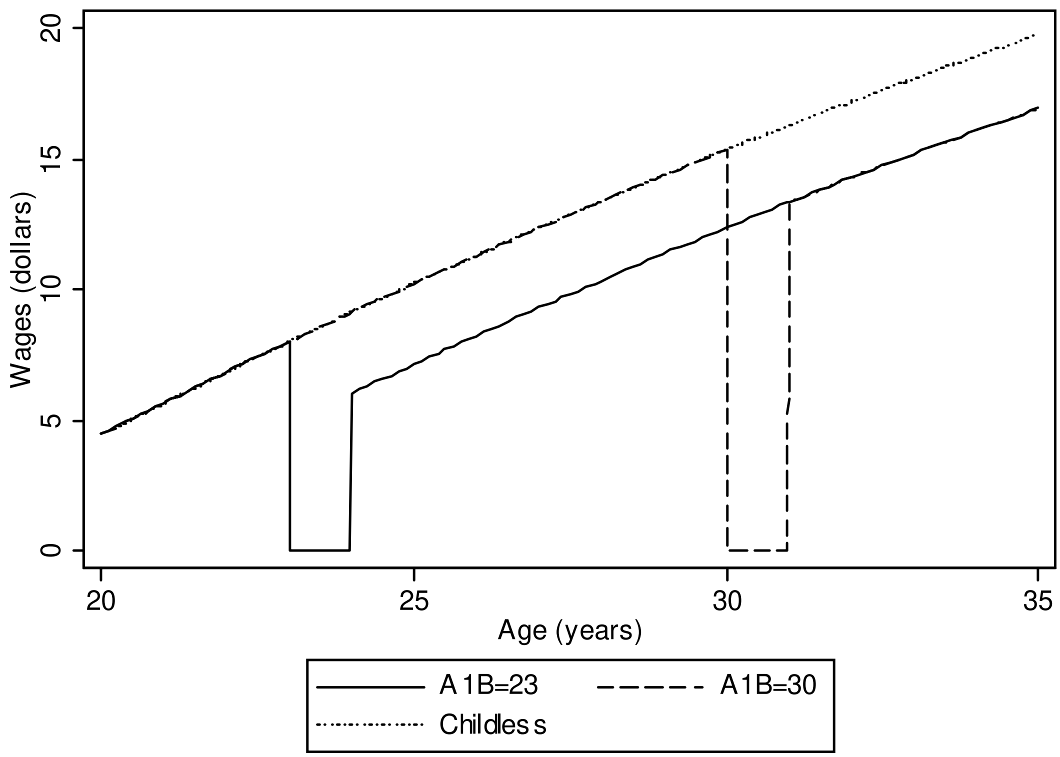
Note: Residuals are the difference between actual and predicted log-wage rate from a linear regression with education, race, ability and birth year cohort as controls.

Figure 6: Returns to Experience



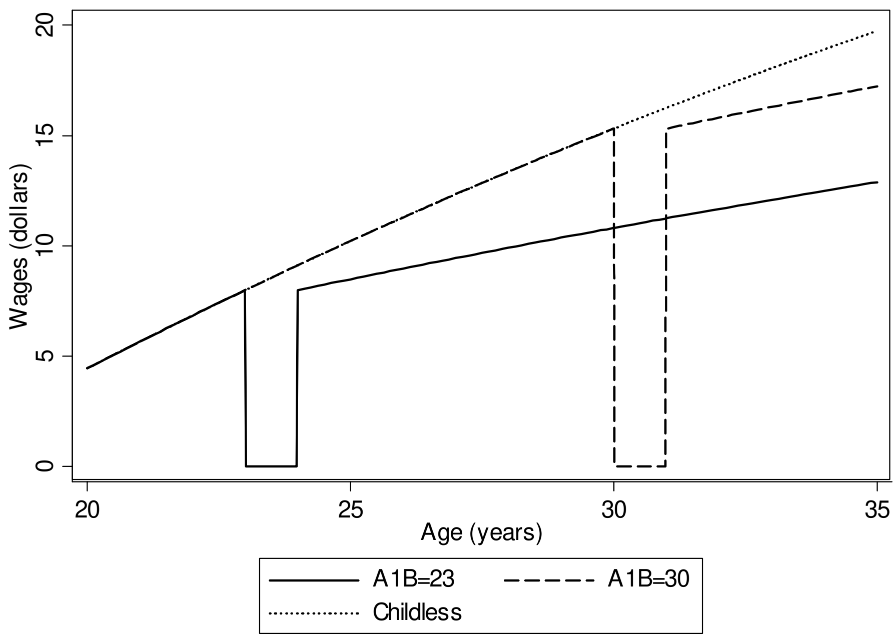
Note: Figure illustrates the effect of a one-year career interruption associated with motherhood at ages 23 and 30, relative to the uninterrupted profile, for a hypothetical woman in Case 3: returns to experience and no additional costs associated with motherhood.

Figure 7: Fixed Costs of Motherhood



Note: Figure illustrates the effect of a one-year career interruption associated with motherhood at ages 23 and 30, relative to the uninterrupted profile, for a hypothetical woman in Case 4: returns to experience and a fixed cost of motherhood or depreciation of human capital.

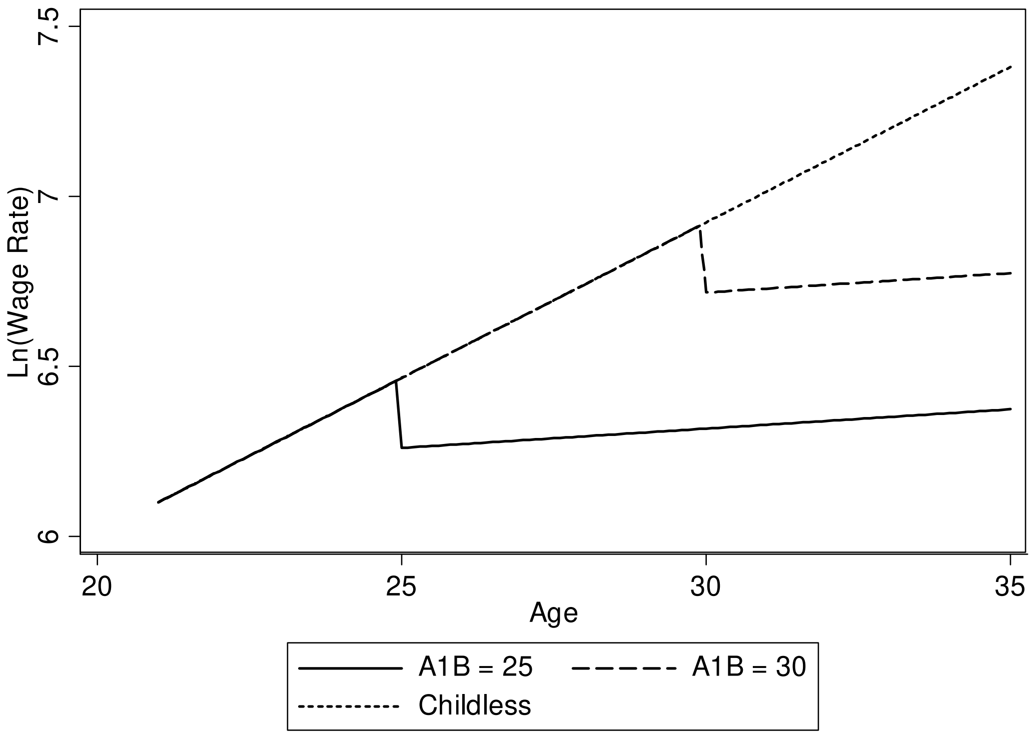
Figure 8: The "Mommy Track"



Note: Figure illustrates the effect of a one-year career interruption associated with motherhood at ages 23 and 30, relative to the uninterrupted profile, for a hypothetical woman in Case 5: returns to experience, with a reduction in returns to experience for mothers.

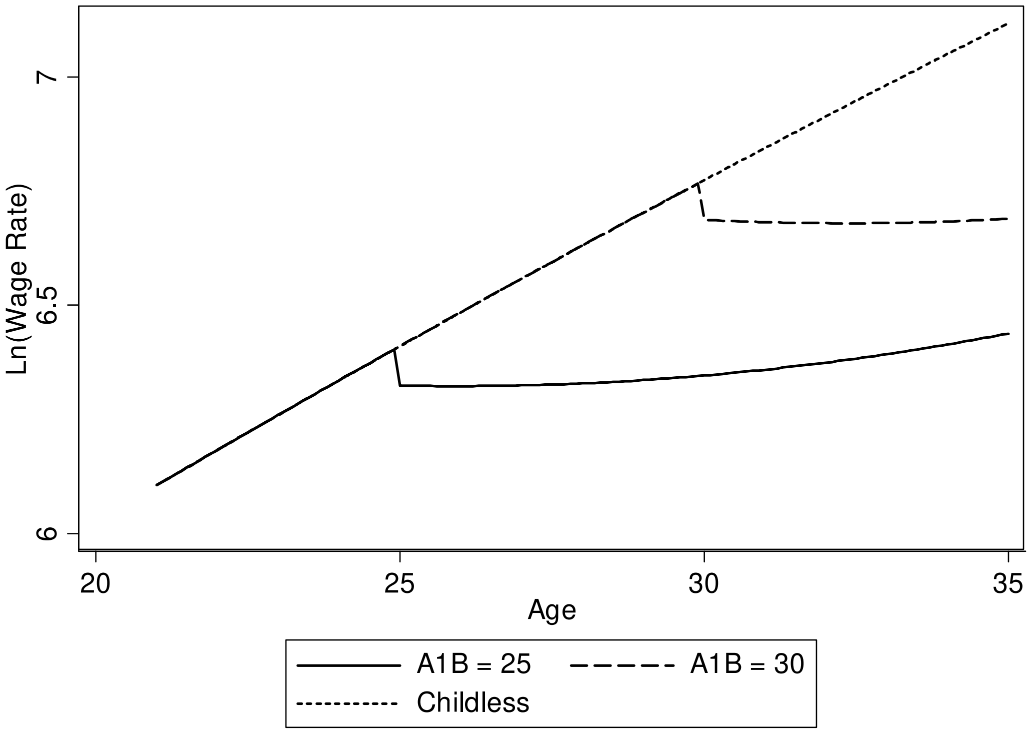
Figure 9: Predicted Age-Wage Profiles by Age at First Birth

Figure 9A – Linear Terms



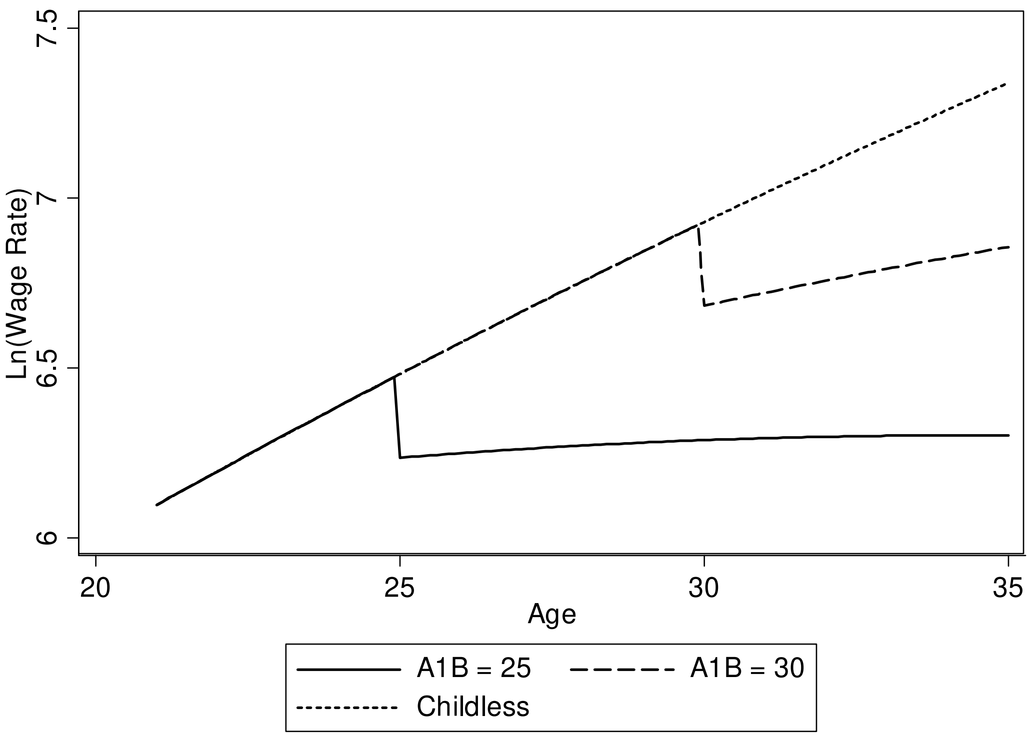
Note: Simulated wage profile using IV regression results reported in Table 6, column 4. The figure represents a set of potential age-wage profiles, for a given woman, depending on her motherhood timing.

Figure 9B – Linear and Quadratic Terms



Note: Simulated wage profile using IV regression results reported in Table 6, column 5. The figure represents a set of potential age-wage profiles, for a given woman, depending on her motherhood timing.

Figure 9C – 3 A1B Categories



Note: Simulated wage profile using IV regression results reported in Table 6, column 6. The figure represents a set of potential age-wage profiles, for a given woman, depending on her motherhood timing.