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Fertility timing, wages, and human capital*

McKinley L. Blackburn[1], David E. Bloom[2], and David Neumark[3]

[1] Department of Economics, University of South Carolina
[2] Department of Economics, Columbia University
[3] Department of Economics, University of Pennsylvania and NBER, 3718 Locust Walk, Philadelphia, PA 19104-6297, USA

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Abstract. Women who have first births relatively late in life earn higher wages. This papers offers an explanation of this fact based on a simple life-cycle model of human capital investment and timing of first birth. The model yields conditions (that are plausibly satisfied) under which late childbearers will tend to invest more heavily in human capital than early childbearers. The empirical analysis finds results consistent with the higher wages of late childbearers arising primarily through greater measurable human capital investment.

I. Introduction

Between 1970 and 1983, the first birth rate in the United States declined by 19%. In addition, according to survey data collected by the Census Bureau, the proportion of childless women increased substantially between 1976 and 1985 in the age groups 25–29, 30–34, and 35–39 (see Table 1). An extensive body of previous research has established that these trends reflect both (1) an increased tendency to permanently forego childbearing and (2) an increasing tendency to delay the initiation of childbearing, among those women who do ultimately bear children (see Bloom 1982, 1984, 1987; Bloom and Pebley 1982; Bloom and Trussell 1984; O'Connell 1985; Rindfuss et al. 1988). In addition, after many years of remarkable stability, the female/male wage ratio also began to increase in the 1980s (from 0.59 in 1980 to 0.67 in 1987). There is some evidence that this latter change is associated with increasing human capital accumulation on the part of women (see Smith and Ward 1984; O'Neill 1985).

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Table 1. Percent childless women, 1976–1985 (by age)

	Year				
	1976	1980	1982	1984	1985
Age group					
18–24	69.0	70.0	72.2	71.4	71.4
25–29	30.8	36.8	38.8	39.9	41.5
30–34	15.6	19.8	22.5	23.5	26.2
35–39	10.5	12.1	14.4	15.4	16.7
40–44	10.2	10.1	11.0	11.1	11.4

Source: US Bureau of the Census, Current Population Reports, Series P-20, No. 406: Fertility of American women (June 1985). US Government Printing Office, Washington DC, 1986

This paper develops some explicit theoretical linkages concerning the relationship between women's fertility-timing behavior and their human capital acquisition and wage profiles. By simplifying the nature of the fertility/work decision, we are able to describe the conditions under which women who prefer (and therefore expect) to be delayed childbearers will invest more heavily in their human capital than women who prefer to have an early first birth. The model suggests that these conditions are fairly general: if the discount rate is greater than the economy-wide growth rate of wages for workers who are not human capital investors, then delayers will be more likely to invest in human capital. Our empirical analysis provides some support for this theoretical model. There exists a strong positive relationship between several proxies for human capital investment and the age at which women bear their first child. However, the empirical analysis does not support the proposition that the usual proxies for human capital represent more intensive investment by women who choose to delay their childbearing.

II. Stylized facts

Our empirical analysis uses the National Longitudinal Survey of Young Women (NLS). This NLS survey has been conducted every year (or every other year in some periods) since 1968, when it started with 5,159 women aged 14–24. The main purpose of this survey has been to gather information on the labor market experiences of young women. Questions on ages of children in various years of the survey enabled us to calculate age at first birth. We have used the NLS data through 1982 to construct a data set with wage, labor market, and fertility timing data for a sample of working women aged 28–38 in 1982. Like many longitudinal data sets, the NLS Young Women's Cohort has suffered from sample attrition over time; still, the 1982 reinterview includes roughly 70% of the original respondents. The NLS data offer a number of advantages compared to other data sets used to study the correlates of fertility timing. For one, the longitudinal structure provides a method of accounting for unobserved heterogeneity. More importantly, perhaps, the richness of the NLS data permits us to construct typically unavailable measures of several key determinants of earnings. The variables used in this study whose construction relies on data from a large section of women's work lives (in many cases their entire career to date) include actual labor market experience,[1] tenure, and occupational training.

Table 2 reports statistics descriptive of wages, human capital, and other characteristics of white women[2] classified into four groups on the basis of the age at which they bore their first child: (1) women who had their first birth before the survey in which they were age 22 ("early" childbearers); (2) women who had their first birth between ages 22 and 26 (inclusive); (3) women who first gave birth after age 26 ("late" childbearers); and (4) women who had not given birth by the time of the 1982 survey ("childless" women).[3] Only individuals with no missing data for any variables (except the training, occupation, and early wage variables) are used to compute the statistics reported in Table 2.[4] The statistics are reported separately for women who were currently employed on the data surveyed in 1982, and for women who were not employed at that time; the differences in the reported statistics across age-at-first-birth categories tend to be very similar for these two samples. With the exception of the early wage variable, all values are for 1982.

Table 2 reveals a cross-sectional pattern of increases in the 1982 wage, and in all of the human capital measures, as age at first birth increases; childless women tend to have values of these variables that are close to those of the late childbearers. These are raw means, of course, and therefore do not control for the influence of some of the variables in the table on others. For example, an explanation for the positive association between wages and age at first birth may be that educational attainment is positively related to age at first birth.[5] Also, the small experience and tenure differences in Table 2 actually do represent a substantially more intensive rate of investment in experience- and tenure-related human capital for delayed childbearers, since their greater investment in education leaves them with relatively less potential time for accumulating experience and tenure.[6] The training measure does not follow as consistent a pattern as education, tenure, and experience, but it does reveal a weak tendency to increase with age at first birth. Finally, later childbearers and childless women are more likely to be in professional and managerial occupations than earlier childbearers, who are more likely to be in administrative, service and blue collar occupations.

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1 This is preferable to potential experience, usually approximated as age minus education minus 5, which is known to be a poor indicator of actual experience for women (see Garvey and Reimers 1980)

2 The sample excludes all nonwhites. Initial investigations suggested that the estimated effect of age at first birth on wages in wage regressions for blacks was very different from the estimated effect for whites. Also, the percentage of blacks who delayed childbearing past the age of 26 was very small.

3 Many of the women who are classified as childless in 1982 will eventually become late childbearers. We would like to separate these women from those who will remain permanently childless, but cannot do so with our data.

4 The training variable is missing for some individuals, and otherwise seemed to contribute little independent information. Consequently, we retained observations for which this variable could not be constructed.

5 While it is not noted in Table 2, it is also true that later childbearers are not delaying simply because they spend more years in school. In fact, higher education is associated with more delaying net of the extra years spent in school. This follows from calculations that show that the difference between age at first birth and the age at completion of schooling is greater for those women with more education.

6 This is clear from the fact that educational attainment increases with age at first birth, while the average age in each age-at-first-birth category is roughly the same (except for the childless, for whom it is the lowest).

Table 2. Wages and other characteristics of white working women aged 28 – 38, by age at first birth (National longitudinal survey of young women: 1982)[a]

	Currently-working women Age at first birth				Non-working women Age at first birth			
	<22	22 – 26	27 +	Childless	<22	22 – 26	27 +	Childless
Wage[b]	5.75	6.59	8.23	7.93	–	–	–	–
	(0.11)	(0.20)	(0.32)	(0.18)				
Education[c]	11.96	13.55	15.01	14.66	11.20	13.25	14.00	12.71
	(0.08)	(0.13)	(0.21)	(0.12)	(0.12)	(0.16)	(0.26)	(0.50)
Experience[d]	7.00	7.85	8.09	7.77	2.67	4.06	5.95	4.96
	(0.17)	(0.25)	(0.33)	(0.18)	(0.19)	(0.19)	(0.31)	(0.51)
Tenure[e]	4.69	5.27	5.50	6.04	–	–	–	–
	(0.16)	(0.27)	(0.37)	(0.23)				
Age	33.39	33.35	33.06	31.77	33.16	32.46	33.34	32.14
	(0.14)	(0.20)	(0.27)	(0.15)	(0.21)	(0.22)	(0.28)	(0.36)
Number of children	2.31	1.90	1.37	–	2.90	2.28	1.64	–
	(0.04)	(0.05)	(0.05)		(0.07)	(0.06)	(0.07)	
% in occupations[f]								
Manager	0.08	0.07	0.11	0.14	0.03	0.03	0.06	0.07
Professional	0.11	0.29	0.45	0.41	0.05	0.21	0.32	0.15
Administrative	0.48	0.41	0.26	0.28	0.29	0.46	0.40	0.31
Service	0.17	0.13	0.11	0.11	0.25	0.17	0.12	0.22
Blue collar	0.15	0.10	0.06	0.07	0.31	0.12	0.10	0.14
No occupation reported	–	–	–	–	0.07	0.01	0.00	0.11

Early wage[g]	4.32	5.26	5.98	5.66	3.82	5.02	5.78	4.86
	(0.14)	(0.14)	(0.21)	(0.13)	(0.19)	(0.15)	(0.19)	(0.23)
Training[h]	0.15	0.15	0.25	0.17	0.10	0.07	0.13	0.26
	(0.02)	(0.02)	(0.06)	(0.02)	(0.02)	(0.02)	(0.03)	(0.07)
N[i]	467	256	115	372	252	192	104	59

[a] Standard errors of means are reported in parentheses. Sample weights were not used in computing the estimates.
[b] The hourly wage is constructed from reported rates and time units of pay.
[c] Highest grade completed.
[d] Actual experience is constructed from a combination of sample-period and retrospective job history questions.
[e] For each year in the survey, the respondent indicates whether or not she is with the same employer as in the previous survey. This information is combined with occasional questions on when the respondent began her last job to construct a measure of years with the current employer. Tenure is reset to zero when a woman is not working at the time of the survey.
[f] For the non-working women, this is occupation of last job, for those who report an occupation.
[g] First available observation prior to 1982 when woman was childless. Adjusted for inflation and productivity to 1982 base using PCE fixed-weight index and index of nonfarm business productivity. Available for sample of 701 women working for a wage, and 290 women not working for a wage, for whom this observation could be constructed.
[h] Constructed from survey questions on duration of occupational and on-the-job training. Measured in year-equivalents (units of 2000 h). Available for 1133 observations on women working for a wage, and 583 women not working for a wage.
[i] Cell sizes are smaller for early wage and training measures.

The early wage variable reported in Table 2 represents an attempt to measure the hourly rate of pay at which women begin their labor market career following the completion of their schooling. In selecting this variable, we found the earliest wage (in 1968 or later) for each women in a year in which she was not in school, and had not previously had her first birth. We were able to measure such a wage for 991 of the 1,817 woman in our sample.[7] The differences in the average level of the early wage across age-at-first-birth categories are similar in pattern to those for the 1982 wage. However, the differences tend to be somewhat larger with the 1982 wage. For example, the average early wage of "late" childbearers is 37% higher than the average early wage for "early" childbearers, while the 1982 wage for late childbearers is 43% higher than the 1982 wage for early childbearers. For the childless compared to the early childbearers, the difference grows from 31% in the early wage to 38% in the 1982 wage.[8]

Some of the relationships revealed in these descriptive statistics replicate results in other studies. Previous analyses of the covariates of age at first birth uniformly reveal that educational attainment varies positively and strongly with first birth timing and the likelihood of being childless (Bloom 1982, 1984; Bloom and Trussell 1984). A related literature documents and studies the relationship between teen childbearing and schooling (e.g., Furstenberg 1976; Furstenberg et al. 1987; Geronimus and Korenman 1992; Hofferth 1984; McCrate 1989; Upchurch and McCarthy 1990). However, there is considerably less research on the direct relationship between women's wages or earnings and the timing of their childbearing.[9] Bloom (1987) finds that in the younger of two cohorts analyzed in the June 1985 CPS, there is a positive relationship between age at first birth and wages, even controlling for schooling, time out of the labor force, and other determinants of earnings. While much has been written about the impact on wages of interruptions in labor force attachment for childbearing and childrearing, this research has not focused on the timing of these interruptions (e.g., Polachek 1975; Mincer and Ofek 1982; Corcoran et al. 1983; O'Neill 1985).

One line of research that does explore specific links between wages and fertility timing is the "business cycle" work of Butz and Ward (1979). The idea underlying this work is that working women will tend to have children when real wages are relatively low, as during a recession. This theory could also have implications for the cross-section relationship between fertility timing and wages, although there have been few attempts to test these implications.[10]

What appears to be lacking in the existing literature is a model that explains the key relationships in Table 2 in a reasonably unified manner. In the next section we offer a theoretical model in which human capital and fertility timing decisions

[7] Since this variable is from different years for different women, we make corrections for differences in the price level and the level of labor productivity across the years.

[8] These comparisons are made using the early wage averages for the part of the sample that was working in 1982.

[9] To our knowledge, the only prior empirical studies of this subject are Trussell and Abowd (1980), Bloom (1987), and Lundberg and Plotnick (1989). Both Lundberg and Plotnick, and Trussell and Abowd, focus on the effects of teenage childbearing only.

[10] Macunovich and Lillard (1989) attempt to apply the Butz and Ward model to micro-data. In contrast to fertility responses to aggregate wage movements, however, they identify effects of wages on fertility timing from cross-section wage variation and time-series wage changes in a panel data set. It seems that a complete test of the Butz and Ward model using micro-data should distinguish expected from unexpected wage changes, and examine the effects of unexpected changes.

are made jointly. Our model suggests that women who prefer to delay the initiation of childbearing will invest more in human capital than women who prefer to begin their childbearing earlier. After presenting this model, Sects. IV and V explore more fully some of its empirical implications.

III. Theoretical model

The key defining feature of the standard human capital model is that individuals have the opportunity to invest in training that enhances their productivity but that is costly to obtain. Workers find it desirable to obtain training when its benefits – higher wages after the training is completed – outweigh its costs. In this section, we extend a simple human capital model to allow childbearing to affect the human-capital investment decision through the effect that withdrawal from the labor force after childbearing has on the benefits of the human capital investment. The key implication of the model is that relatively late childbearing is associated with greater human capital investment; factors that lead to relatively late childbearing also lead to greater human capital investment, and vice versa.

In analyzing the human-capital/fertility-timing decision for women, we make several simplifying assumptions.[11] First, we assume that all women are equally productive in the labor market at the start of their working career. Second, all women bear one, and only one, child. Third, all women work from time 0 to time R, except for a period of length τ following childbirth when women leave the work force and have no earnings; women must choose to have their child at some time between time 0 and time $R-\tau$. Fourth, all women have the option of investing in one type of human capital which increases a woman's productivity (and thus her wage), at the rate s, while she is working; the cost of the human capital investment is C (at time 0) and is the same for all women. Finally, the only source of income for a woman is her own earnings.

We also assume that a woman's lifetime utility depends on the present value of her lifetime income and on the time period in which she has her child, i.e., $U=U(Y,T)$ where Y is the present value of lifetime income and T is the time period of childbirth; women prefer both higher incomes and earlier periods of childbearing (i.e., we assume $U_Y>0$ and $U_T<0$). For tractability, we use a linear utility function, $U=Y-aT$, with $a>0$.[12] Women differ in their preferences toward early childbearing, so that a varies across women.

At the start of her working career, a woman has two choices to make: first, whether she should invest in human capital; and, second, when she should have her first child. We seek a description of these two choices in the following way: first, we derive an expression for the optimal age at first birth conditional on either investment (with optimal age T_I^*) or non investment (T_N^*); second, given that the woman is optimizing in her choice of time of childbirth, we ask whether her lifetime utility would be higher under investment or under non-investment.

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11 A related but simpler model, due to Edlefsen (1980), is presented in Montgomery and Trussell (1986). Happel et al. (1984), Razin (1980) and Cigno and Ermisch (1989) also present economic models of fertility timing decisions, but only Cigno and Ermisch have treated human capital accumulation as determined jointly with fertility timing.

12 The important aspect of this assumption for the analysis of this paper is that utility be strongly separable in the woman's labor income and fertility timing.

The complete details of the solution for these two decisions are quite cumbersome (due to the possibility of boundary solutions for time of birth) and therefore are presented in the appendix. Here we highlight the nature of the solution by focusing on the case where all women choose an interior solution for time of birth (i.e., $0<T^*<R-\tau$). If a woman chooses to invest in human capital, then the present value of her lifetime income is

$$Y_I=\int_0^T w\mathrm{e}^{(s+q-r)t}\mathrm{d}t+\int_{T+\tau}^R w\mathrm{e}^{(q-r)t}\mathrm{e}^{s(t-\tau)}\mathrm{d}t-C ,$$

where w is the initial wage, q is the economy-wide rate of growth of wages, and r is the discount rate. $U(Y_I, T)$ is maximized with respect to T, for which the first-order condition for a utility extremum is

$$\mathrm{MB}_I(T)\equiv w\mathrm{e}^{(s+q-r)T}[1-\mathrm{e}^{(q-r)\tau}]=a . \tag{1}$$

The left-hand side of (1) represents the marginal benefit from delaying childbirth, and equals the present value of the wage received prior to childbirth minus the present value of the wage received when returning to the labor market after childbirth. The right-hand side of (1) is the marginal cost of delaying. If follows that the marginal benefit of delaying will only be positive if $q<r$, which we assume throughout. The condition will represent a maximum if $s+q-r<0$, which we also presently assume. (If $s+q-r>0$, the optimal time for the birth will be either 0 or $R-\tau$; this case is discussed in the appendix).

If the woman chooses not to invest in human capital, lifetime income is

$$Y_N=\int_0^T w\mathrm{e}^{(q-r)t}\mathrm{d}t+\int_{T+\tau}^R w\mathrm{e}^{(q-r)t}\mathrm{d}t .$$

The condition for a utility extremum is

$$\mathrm{MB}_N(T)\equiv w\mathrm{e}^{(q-r)T}[1-\mathrm{e}^{(q-r)\tau}]=a , \tag{2}$$

which is also the condition for a maximum when $q<r$. The left-hand side of (2) represents the marginal benefit of delaying for a non-investor.

It follows from (1) and (2) that a women who has invested in human capital will choose to delay her childbearing more than if she had not invested in human capital. This fact is illustrated in Fig. 1, which graphs the marginal benefit and marginal cost functions for both investors and non-investors. While the two marginal benefit functions are equal when $T=0$, at any time T the slope of the investor's marginal benefit function is larger (less negative) than the slope of the noninvestor's function. As a result, $T_I^*>T_N^*$ must hold. As a decreases (as in the fall from a_0 to a_1 in Fig. 2), the larger slope of the investor's marginal benefit function implies that the difference between T_I^* and T_N^* will increase, as long as T_I^* has not reached $R-\tau$, the upper boundary for childbearing ages.[13]

[13] Note also that as s increases, the investor's marginal benefit function will grow less steep (swinging out to the right but keeping the same y-intercept), which increases $T_I{}^*$, and the difference between $T_I{}^*$ and $T_N{}^*$.

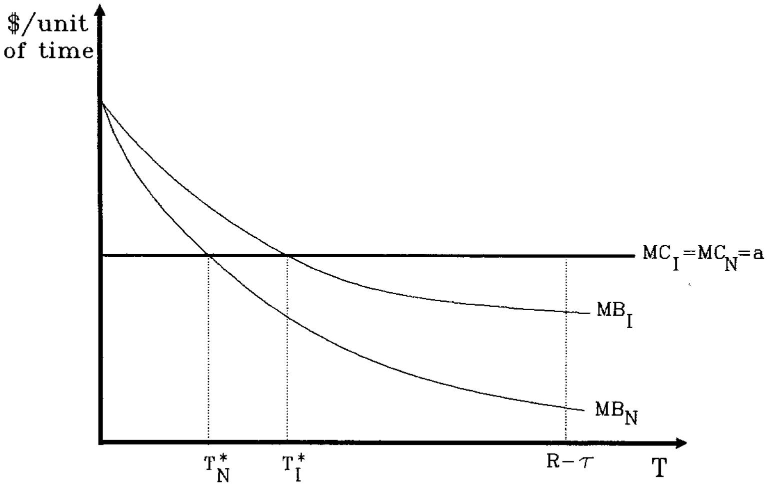

Fig. 1

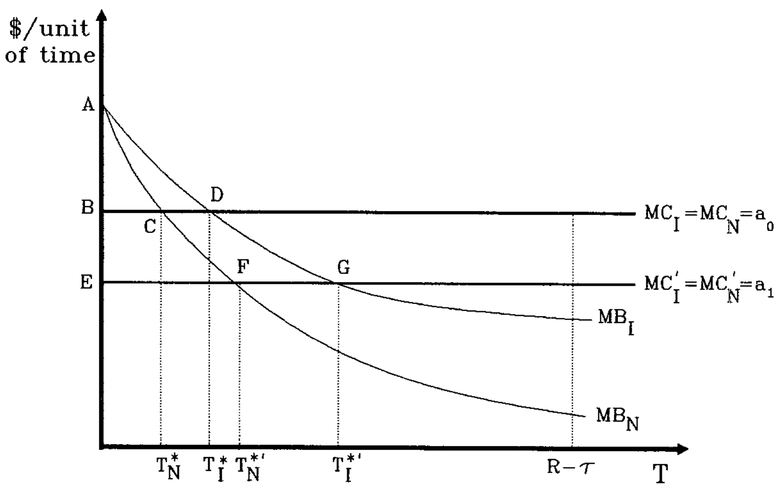

Fig. 2

The next step in our analysis of the decision process is the human capital investment decision. Given that the woman will be optimizing on her fertility timing once she has chosen whether or not to invest, we can derive the indirect utilities of the two choices as functions of the parameters of the problem, i.e., $V_j = V_j(w, q, r, a, s, C)$, $j = I, N$. Then a women will choose to invest in human capital if $V_I > V_N$. As is demonstrated in the appendix, the indirect utility from investing grows more quickly as a declines than does the indirect utility from not investing; this means that women with weaker preferences for early childbearing are more likely to find $V_I > V_N$ to be true than women with stronger preferences for an early birth.

This result is illustrated in Fig. 2 for the case where both investors and non-investors choose an interior solution for T^*. Define the surplus from delaying childbearing as the total additional benefits associated with delaying childbearing past $T=0$ minus the total additional (utility) costs. For instance, with childbearing preference parameter $a=a_0$, the surplus from delaying for non-investors is $S_N(a_0)=\text{area}(ABC)$ in Fig. 2, and the surplus for investors is $S_I(a_0)=\text{area}(ABD)$. While it is obvious that $S_I(a_0)>S_N(a_0)$, a woman with utility parameter a_0 will choose to invest only if

$$S_I(a_0)-S_N(a_0)>-(V_I-V_N)|_{T^*=0}\ , \tag{3}$$

i.e., if the additional surplus associated with investing is greater than minus the difference in indirect utility between investors and non-investors if both were to have their child at time 0. $(V_I-V_N)|_{T^*=0}$ does not depend on the value of a (see case 1 of the Appendix).[14]

Consider another woman with utility parameter $a=a_1$, where $a_1<a_0$. Her surplus from delaying will be $S_N(a_1)=\text{area}(AEF)$ if she does not invest, and $S_I(a_1)=\text{area}(AEG)$ if she does. Going from a_0 to a_1, the change in both surpluses will be positive, i.e. $\Delta S_I=S_I(a_1)-S_I(a_0)=\text{area}(BEGD)$ and $\Delta S_N=S_N(a_1)-S_N(a_0)=\text{area}(BEFC)$ are both greater than zero. But as is clear from Fig. 2, $\Delta S_I>\Delta S_N$, implying that it is more likely for a woman with the lower value of a to find investing optimal than for a woman with the higher value, since the left-hand side of (3) increases as a decreases but the right-hand side remains unchanged.

What does this imply for the earnings of women who choose to delay childbearing? As discussed earlier, we expect women with lower values for a to delay their childbearing more, while we also except them to be more likely to invest in human capital. With human capital investment increasing the wages of investors relative to non-investors at all times after $T=0$, we expect to observe delayed childbearers having higher wages than otherwise similar women (in terms of initial productivity) who choose to bear their child earlier. It also follows that the difference in the wages of delayed and early childbearers will grow over time due to the higher rate of human capital investment among delayed childbearers.

While the model has been discussed in terms of variation in the taste parameter a being the only difference between women, it is also true that increases in s, or decreases in C, make it more likely both that women will delay childbearing and that they will invest in human capital. Furthermore, slight changes in the model do not alter the basic conclusions. For example, if the model is changed to reflect the fact that women have more difficulty continuing their investment in human capital after having their first child, so that wages grow at the rate q (and not at the rate $q+s$) after returning from childbirth even for women who were earlier investing in human capital, our primary conclusions still hold. Likewise, if there is depreciation of human capital while a woman is out of the labor force, delayers will still invest more in human capital, as depreciation merely increases the difference between the leaving and returning wages and so "swings out" the $\text{MB}_I(T)$ curve.

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[14] This difference may be positive, though in this case all women would end up choosing to make the human capital investment.

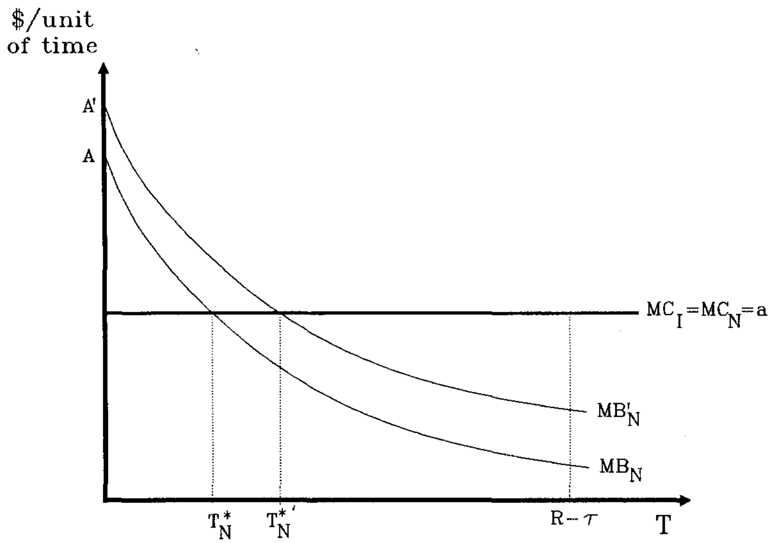

Fig. 3

Differential investment in human capital is not the only reason why delayed childbearers might be expected to have higher wages. Suppose there were no possibility of investing in human capital, and that all women had the same value for the childbearing utility parameter *a*. Instead, women differ in their initial level of productivity (and so their wage) at the time they enter the labor force. From (2) we know that a higher initial wage will shift upwards the marginal benefit curve, as from MB_N to MB'_N in Fig. 3. As a result, women with higher initial wages will more often choose to delay, and so delayed childbearing will be associated with higher wages at all points in the woman's lifetime. Differences in initial wages may also be part of the explanation for the fact, discussed in Sect. II, that more educated women tend to delay their childbearing more after completing their schooling. [15]

The empirical implications of our model are that women who choose to delay childbearing will accumulate more education, and, to the extent that labor market experience and tenure represent human capital investment, will also accumulate more experience and tenure. This greater investment will raise the relative earnings of later childbearers. We might expect two additional findings, because we may not be able to perfectly measure human capital investment, but rather must regard variables such as education, experience, tenure, and even our training measure, as proxies. First, we may find that fertility timing has an effect on wages after controlling for measurable human capital, though such residual effects could also be due to unobserved heterogeneity in the initial productivity of workers. Second, we may find higher returns to education, experience, and tenure, for delayed childbearers, in wage regressions. We explore these possibilities in the next section.

[15] Cigno and Ermisch (1989) suggest that post-marital investments in human capital will be larger for delayed childbearers. In their model, higher pre-marital investments in human capital increase a woman's tempo of fertility, but reduce completed family size.

IV. Wage equation estimates and fertility timing

It is difficult if not impossible to measure human capital directly. As a result, labor economists often interpret wage variation associated with variation in certain observable factors as resulting from differences in investment in human capital, relying on a theoretical structure that relates the observable variables to human capital investment. One famous example is Mincer's interpretation of labor market experience as representing investments in on-the-job training (see Mincer 1974). In the present context, this approach suggests including age-at-first-birth variables in a wage regression as potentially reflecting differences in human capital investment not captured by other human capital proxies that are included in the regression. In this section, we consider this approach, as well as more explicit tests of the human capital model.

Table 3 presents least squares estimates of log wage equations for the NLS sample of working women considered in Table 2. The wage variable is the report-

Table 3. Wage equation estimates for white working women, ordinary least squares (Dependent variable: natural logarithm of hourly earnings)[a]

	(1)	(2)	(3)	(4)
Age at first birth				
22–26	0.10	0.01	−0.03	−0.03
	(0.03)	(0.03)	(0.03)	(0.03)
27+	0.32	0.14	0.07	0.06
	(0.05)	(0.05)	(0.04)	(0.04)
Childless	0.30	0.14	0.07	0.06
	(0.04)	(0.04)	(0.04)	(0.04)
Joint significance (p-value)[b]	0.00	0.00	0.05	0.06
Years of education	–	0.06	0.07	0.06
		(0.01)	(0.01)	(0.01)
Post-college dummy variable	–	−0.07	−0.05	−0.06
		(0.04)	(0.04)	(0.04)
Experience	–	–	0.02	0.02
			(0.003)	(0.003)
Tenure	–	–	0.05	0.05
			(0.01)	(0.01)
Tenure[b] $\times 10^{-2}$	–	–	−0.16	−0.15
			(0.05)	(0.04)
R^2	0.16	0.23	0.34	0.36
Occupation dummy variables included[c]	No	No	No	Yes

[a] There are 1210 observations. Standard errors of estimates are reported in parentheses. Controls included in all specifications include: dummy variables for two children and three or more children; dummy variables for marital status (married, spouse present and divorced, widowed or separated); and dummy variables for residence in the South and in an SMSA. Regressions in columns (1) and (2) also include age entered as a linear variable; the estimated coefficient of a quadratic term was statistically insignificant, as were the estimated coefficients of the linear and quadratic terms in columns (3) and (4). The estimated coefficient of a quadratic term in experience was statistically insignificant in columns (3) and (4), and was omitted.

[b] Computed from standard F-test.

[c] The occupation categories in Table 2 are used to define dummy variables for occupation.

ed hourly wage and salary income usually earned in the respondent's primary job at the time of the survey. In all of our regressions, we include as independent variables dummy variables for the same classification of age at first birth as was used for Table 2; age at first birth less than 22 is the omitted category.[16] In column (1), we report coefficient estimates from a least-squares regression of the natural logarithm of the wage on these timing dummy variables, as well as dummy variables for living in the South and living in an SMSA (crude controls for cost-of-living differentials). In addition, we include as regressors age at the time of the survey, and dummy variables for whether the woman was married with spouse present, or was instead divorced, widowed or living apart from her spouse (never married is the omitted category).[17] Since the number of children to which a woman has given birth is correlated with her age at first birth (see Table 2), we also include dummy variables for two children, and for three or more children, as independent variables. The choice of omitted category for the number of children classification implies that the childless coefficient measures the difference between childless women and early childbearers with one child. As the estimates in Table 3 make clear, substantial wage differentials by age at first birth persist once these controls are added. Women with age at first birth between 22 and 26 earn on average 10% more than early childbearers, and these differentials rise to about 30% for late childbearers and childless women.

Because few (if any) controls for human capital investment are included in the specification in column (1), these estimates of wage differentials due to age at first birth may reflect differences in the observable proxies for human capital such as education and experience. In column (2), we add education to the regression.[18] Not surprisingly, the wage differentials by age at first birth fall. Nonetheless, sizable and statistically significant differentials remain for late childbearers and childless women, consistent with differences in human capital investment above and beyond education.

Human capital investment may also occur on the job. In column (3) we add experience, tenure and its square to the regression.[19] The coefficient estimates for experience and tenure are statistically significant, and of the expected sign and magnitude. The age-at-first-birth coefficients are further reduced by the inclusion of experience and tenure, and though individually are statistically insignificant, remain significant in a joint *F*-test. Inclusion of occupational dummy variables – which may reflect human capital investment differences – further reduces the coefficient for "late" childbearers, leaving the age-at-first-birth dummy coefficients significant only at the 0.06 level.

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[16] In unreported results, we experimented with linear, quadratic, and other specifications of age-at-first-birth effects. A linear specification was clearly inadequate, and the dummy variable specification is more readily interpretable than other non-linear specifications that appeared to fit the data equally well.

[17] Women's marital status has been linked theoretically to differential human capital investment (Becker 1985), though Korenman and Neumark (1992) find some empirical evidence to the contrary.

[18] The education variables are years of schooling, and a dummy for post-college (i.e., more than 16 years of schooling). Specifications including dummy variables for high school and college degrees were also estimated, but the coefficient estimates were statistically insignificant and small.

[19] Coefficient estimates for experience squared were insignificant; excluding this variable has little effect on other coefficient estimates. We also exclude age at survey in specifications that include experience and tenure.

The estimates in Table 3 support two important findings. First, differences in observed proxies for human capital investment (schooling, experience, and tenure) can explain a sizable portion of wage differentials associated with fertility timing. Second, wage differentials remain once account is taken of these proxies, consistent with there being differences in unobserved human capital investment (although the statistical evidence on this point is not strong). We next ask whether these remaining fertility-timing effects actually reflect further differences in human capital investment.

Including dummy variables for fertility timing may be a crude way to estimate effects of differential human capital investment that remain once our proxies are included. A better specification may be to let the coefficients of these proxies vary with age at first birth. These types of effects could arise if the education of women who intend to delay childbearing is associated with more intensive human capital investment per year of schooling (thus giving rise to higher returns to education), or if later childbearers invest more per unit of labor market experience or tenure. In column (1) of Table 4, we let the coefficient on the years of schooling variable differ for late childbearers and childless women by including interactions

Table 4. Wage equation estimates for white working women, interactive specifications and incorporating training measures, ordinary least squares (Dependent variable: natural logarithm of hourly earnings)[a]

	Interactive specifications			Training specifications	
	(1)	(2)	(3)	(4)	(5)
Age at first birth					
22 – 26	−0.02	−0.03	−0.02	−0.03	−0.03
	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)
27 +	0.04	0.07	0.05	0.08	0.07
	(0.05)	(0.04)	(0.05)	(0.04)	(0.04)
Childless	0.07	0.07	0.07	0.06	0.06
	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)
Joint significance (p-value)[b]	0.17	0.04	0.12	0.06	0.08
Years of education	0.05	0.05	0.05	0.06	0.06
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Post-college dummy variable	−0.08	−0.06	−0.08	−0.04	−0.04
	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)
Experience	0.02	0.02	0.02	0.02	0.02
	(0.003)	(0.004)	(0.004)	(0.003)	(0.003)
Tenure	0.05	0.05	0.05	0.05	0.05
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Tenure[b] $\times 10^{-2}$	−0.15	−0.13	−0.13	−0.17	−0.17
	(0.04)	(0.05)	(0.05)	(0.05)	(0.05)
Training	–	–	–	–	0.05
					(0.03)
Age at first birth x education					
27 +	0.029	–	0.027	–	–
	(0.017)		(0.019)		
Childless	0.014	–	0.011	–	–
	(0.011)		(0.012)		
Joint significance (p-value)[b]	0.17	–	0.31	–	–

Table 4 (continued)

	Interactive specifications			Training specifications	
	(1)	(2)	(3)	(4)	(5)
Age at first birth x experience					
27+	–	−0.014	−0.007	–	–
		(0.011)	(0.012)		
Childless	–	−0.004	−0.003	–	–
		(0.007)	(0.008)		
Joint significance (p-value)[b]	–	0.41	0.83		
Age at first birth x tenure					
27+	–	0.013	0.013	–	–
		(0.010)	(0.010)		
Childless	–	−0.010	−0.010	–	–
		(0.006)	(0.006)		
Joint significance (p-value)[b]	–	0.06	0.05	–	–
R^2	0.37	0.37	0.37	0.37	0.37
N	1210	1210	1210	1133	1133

[a] Standard errors of estimates are reported in parentheses. Controls included in all specifications include: dummy variables for two children and three or more children; dummy variables for marital status (married, spouse present and divorced, widowed or separated); dummy variables for residence in the South and in an SMSA; and occupation dummy variables. The education, experience, and tenure interactions are products of the age at first birth dummy variables and the respective variable minus its sample mean; thus, for example, the coefficient on years of education continues to measure the effect of education on log wages evaluated at the sample means.
[b] Computed from standard F-test.

of education with the 27+ and childless dummies.[20] In column (2), we let the linear experience and tenure coefficients vary by age-at-first-birth category, while in column (3) we let the returns to all three human capital proxies vary. Focusing on column (3), we see that the point estimates of the education interactions are consistent with greater human capital investment for each year of education for later childbearers. For example, the return to education is almost 60% higher for the late childbearers (27+) relative to the <22 category.[21] But these differences are not statistically significant. While the point estimates for the experience interactions are small and insignificant, the tenure interactions are jointly significant. The tenure coefficient estimates suggest that late childbearers receive a higher return to tenure, but also provide the anomalous result that childless women receive lower returns to tenure (this latter inference supported by a marginally significant t-statistic). Overall, the patterns in these interactions do not support the proposition that our human capital measures represent more human capital investment per unit of education and experience for women who delay their childbearing than for early childbearers.

[20] An interaction for the 22–26 category was not included because the estimates in Table 3 provided no evidence that these women earn higher wages than early childbearers once observable human capital proxies are included in the wage regression.

[21] In unreported results, we verified that these differences do not simply reflect nonlinearities in education effects entailing higher returns to the higher levels of education of later childbearers.

As a final attempt to better capture human capital differences, we utilize the training measure presented in Table 2 as a more direct measure of human capital investment. In column (4) of Table 4 we repeat the estimation of the specification in column (4) of Table 3 using the smaller sample with nonmissing training. Then, in column (5), we add to this specification the training measure. The estimates suggest that each additional year of training leads to a 5% increase in wages. However, inclusion of training does not affect the fertility timing coefficient estimates, or the coefficient estimates for experience and tenure.[22]

Our next step in the empirical work is to explore whether these conclusions are affected by the failure to control for two other potential influences on the relationship between wages and fertility timing: the effect of fertility timing on labor force participation; and the direct effect of wages on fertility timing.

Since the estimates in Tables 3 and 4 are based on samples of women that exclude those women not working in 1982, inferences from these estimates may suffer from sample selection bias (see Heckman 1979). The nature of our sample definition implies that our estimates are conditional on women working, and so may not correspond to the population of all women. In particular, it seems plausible that the age-at-first-birth coefficients may be affected by our sample selection, since from Table 2 there are clear differences in the proportion of women working across age-at-first-birth categories. (This proportion varies from 53% among the late childbearers to 86% among the childless.) This difference may be due to reservation wages varying systematically by womens' ages at first birth, so that late childbearers have higher reservation wages, and childless women lower reservation wages, than earlier childbearers.[23] If so, then we would expect equations estimated with the complete population (i.e., not conditional on working in 1982) to exhibit larger differences in wages between early childbearers and the childless, and smaller differences between the early childbearers and the later childbearers, than we observed in earlier tables.

To study the influence of selectivity on our wage equation estimates, we reestimated specification (2) in Table 3 – where education, but not experience and tenure are included – using a maximum-likelihood procedure for the two-equation model suggested in Heckman (1979).[24] In column (1) of Table 5, we present estimates from a model where several variables are included as determinants of working status (husband's income and unemployment, alimony and child support, and several family background variables) but are assumed not to belong in the wage equation specification. Compared to Table 3, these results suggest that – consistent with our expectations – the 27+ effect was overstated in our earlier estimates, while the childless effect was understated; but the changes in the coefficients due to the selection correction are not very large. Since it is also possible to correct for selectivity without imposing the restriction that the additional variables (mentioned above) be excluded form the wage equation, we re-estimated our

22 One interpretation of these results is that much of the human capital investment undertaken by delayed childbearers does not occur in the formal settings captured by the training variable.

23 Reservation wages would follow this pattern if the presence of children (and especially young children) raised the opportunity cost of market time for women.

24 Experience and tenure were not included in the wage equation or the probit equation for working in our selection-corrected estimations. Tenure could not be included, since it is zero if and only if the woman is not working, while past labor market experience is very likely to be correlated with the error term in the probit equation for currently working.

selection model without imposing these restrictions.[25] These alternative estimates are presented in column (2). In contrast to the previous results, with this specification the 27+ coefficient estimate is unchanged while the childless coefficient drops, relative to the least-squares estimates. However, there are two reasons to prefer the estimates in column (1): first, the likelihood-ratio statistic ($\chi^2 = 14$, with 13 degrees of freedom) for testing the exclusion restrictions in column (1) is not significant; and, second, the positive sign of the error correlation in column (1) is more plausible.[26]

The idea that wages may directly affect fertility timing was illustrated in the discussion of Fig. 3 in Sect. III. This discussion suggested that fixed omitted variables in a wage equation (e.g., ability, aggressiveness, and "spunk"), which would tend to increase starting wages, should also lead to delayed childbearing.[27] We consider this possibility in two ways. First, we attempt to control for the fixed wage effect that may be correlated with age at first birth in estimating our wage equations for 1982. We do this by including as a regressor the residual from a regression of the early (log) wage on characteristics of women at the time of the early wage observation.[28] In column (4) of Table 5, we re-estimate the specification reported in column (4) of Table 3 using the smaller sample for which the early wage variable was available. In column (5) of Table 5, we add the early wage residual.[29] The results suggest that there is some persistence in wages over time, as the coefficient estimate for the early wage residual (which is measured, on average, 11 years prior to 1982) is 0.23. The coefficient estimates for the fertility timing dummy variables fall when the early wage residual is included, although the decline in the coefficients is small relative to the standard errors of the coefficient estimates. The implication is that heterogeneity in initial wages does not affect to any great extent the measured relationships between wages and fertility timing. In column (3), we report coefficient estimates for the age-at-first-birth dummies when they are included as regressors in the early wage equation. These estimates also reveal a weak correlation between starting wages and the timing of (later) fertility once the effects of human capital on wages are removed.[30]

Our second method for handling the potential correlation between age at first birth and wages is to instrument for the age-at-first-birth dummies. This method is potentially more useful because it can correct for such a correlation arising

[25] In this specification, identification comes from the nonlinearity associated with the assumption of bivariate normality for the error terms.

[26] The negative correlation coefficient found in specification (2) suggests that low-wage women are more likely to be working than high-wage women, all else the same. Given that exogenous income variables are included in the probit equation, this seems unlikely.

[27] The presence of fixed effects in the wage equation error is a plausible explanation for why there may be a correlation between the wage equation error and the age-at-first-birth variables, leading to biased coefficient estimates. However, since current wages should have little direct effect on past childbearing behavior, a correlation between the current wage and the age-at-first-birth variables is unlikely to arise from a correlation between age at first birth and the current-period innovation in the wage equation error.

[28] We cannot estimate a standard fixed-effects model because the age-at-first-birth dummy variables are fixed over time.

[29] The age-at-first-birth dummies were not included in the early wage regressions used to generate the early wage residuals.

[30] Although not reported in the table, we also estimated the interactive specifications reported in Table 4 correcting for selectivity or heterogeneity; the results of these specification were unaffected by these corrections (see Blackburn et al. 1990).

Table 5. Wage equation estimates for white working women, biases from selection, heterogeneity, and endogeneity (Dependent variable: natural logarithm of hourly earnings)[a]

	Sample-selectivity correction[b]		Heterogeneity correction[c]			Endogeneity corrections[d]			
			Early	1982	1982	OLS	Endogenous age at first birth	OLS	Endogenous experience and tenure
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Age at first birth									
22–26	−0.01	0.02	0.01	−0.05	−0.06	−0.02	0.20	−0.06	0.03
	(0.03)	(0.04)	(0.04)	(0.05)	(0.05)	(0.05)	(0.92)	(0.05)	(0.07)
27+	0.10	0.14	0.01	0.05	0.04	0.11	0.24	0.03	0.18
	(0.05)	(0.06)	(0.05)	(0.06)	(0.06)	(0.06)	(0.59)	(0.06)	(0.10)
Childless	0.15	0.11	0.02	0.08	0.06	0.11	0.51	0.05	0.16
	(0.04)	(0.05)	(0.04)	(0.06)	(0.06)	(0.05)	(0.64)	(0.06)	(0.08)
Joint significance	0.00	0.01	0.94	0.04	0.05	0.00	0.22	0.07	0.05
Experience	–	–	0.04	0.02	0.02	–	–	0.014	−0.050
			(0.01)	(0.004)	(0.005)			(0.005)	(0.032)
Tenure	–	–	0.05	0.05	0.05	–	–	0.024	0.047
			(0.01)	(0.01)	(0.01)			(0.004)	(0.031)
Early wage residual			–	–	0.23	0.28	0.29	0.22	0.27
					(0.04)	(0.04)	(0.05)	(0.04)	(0.06)
Estimation method	MLE	MLE	OLS	OLS	OLS	OLS	IV	OLS	IV
Corresponding OLS estimates	Table 3, column (2)	Table 3, column (2)	–	–	–	–	Column (6)	–	Column (8)
Rho	0.33	−0.13							
	(0.14)	(0.43)							
Log-likelihood	−1552.4	−1548.5							
Hausman tests (separate regressions)									
Age at first birth 22–26 residual						–	−0.22	–	–
							(0.88)		

Age at first birth 27+ residual	–	−0.14 (0.56)	–	–
Childless residual	–	−0.40 (0.61)	–	–
Joint significance	–	0.51	–	–
Experience residual	–	–	–	0.09 (0.03)
Tenure residual	–	–	–	−0.01 (0.03)

[a]Asymptotic standard errors of estimates are reported in parentheses.

[b]Based on 1817 observations, with 1210 working women. Controls included in wage equation and employment probit include: age; dummy variables for two children and three or more children; dummy variables for marital status (married, spouse present and divorced, widowed or separated); dummy variables for residence in the South and in an SMSA; years of education and a dummy variable for post-college education. In column (2) the following family background variables are included in both equations: husband's income (set to zero for unmarried women); the sum of income from alimony and child support (set to zero for never married women); father's education; mother's education; number of siblings; a dummy variable equal to one if the respondent's mother worked when respondent was age 14; a dummy variable equal to one if the respondent lived with both a father and a mother at age 14; and dummy variables corresponding to each of these variables, equal to one when the variable was missing (in which case the variables were set equal to zero). In column (1) these variables were excluded from the wage equation.

[c]Based on 701 observations. Controls included in columns (3) – (5) include: years of education and a post-college dummy variable; tenure squared; dummy variables for marital status; dummy variables for residence in the South and in an SMSA; and occupation dummy variables. In each case, the years from which these controls are drawn are indicated in the column heading. The age at first birth variables, however, are always as defined in 1982. In column (3) dummy variables for the year from which the early wage are drawn are included. The early wage residual is computed from the early wage regression excluding age at first birth variables.

[d]Based on 701 observations. Controls included in columns (6) and (7) include: age; years of education and a post-college dummy variable; and dummy variables for residence in the South and in an SMSA. Additional controls included in columns (8) and (9) are: dummy variables for two children and three or more children; and dummy variables for marital status. The square of tenure was excluded. Instrumental variables used in column (7) were: father's education; mother's education; number of siblings; a dummy variable equal to one if the respondent's mother worked when respondent was age 14; a dummy variable equal to one if the respondent lived with both a father and a mother at age 14; and dummy variables corresponding to each of these variables, equal to one when the variable was missing (in which case the variables were set equal to zero). Instruments for experience and tenure in column (9) include these variables, as well as: husband's income (set of zero for unmarried women); the sum of income from alimony and child support (set to zero for never married women). Residuals reported in last five rows of table for columns (6) – (9) are from regressions of potentially endogenous variables on all exogenous variables and instruments. These residuals are then included in the OLS regression, and their coefficients (standard errors) are reported. Statistically insignificant coefficient estimates imply that exogeneity cannot be rejected.

from heterogeneity or endogeneity. We estimate linear probability models for each of the three first-birth dummies, using family background variables available in the NLS (as well as the exogenous variables in the wage equation) as independent variables. We then use predicted probabilities from these estimated models as instruments for the first-birth dummies in the 1982 wage equation (see Heckman 1978). As experience and tenure might also be endogenous in the wage equation, we estimate a wage equation that excludes these variables. We also exclude other demographic variables (marital status, and number of children), allowing age at first birth to capture all possible demographic effects. However, we include the early wage residual as a regressor; if this residual captures unobserved ability, it should be valid to exclude family background variables from the wage equation (see Griliches 1979). The results in column (6) of Table 5 are OLS estimates for this equation, while column (7) reports our instrumental variables estimates.[31] The IV estimation increases the coefficient estimates for the first-birth dummies, but also leads to considerably less precise estimates. Results from a Hausman test – performed by including the residuals from the linear-probability models as additional regressors – suggest that we cannot reject the hypothesis that the fertility timing variables are exogenous. (The p-value for the joint significance of these residuals is 0.51).

We also estimated the variant of our wage equation model that includes experience and (linear) tenure, treating these two variables as potentially endogenous in the wage equation. Given the results of the Hausman test for column (7) of Table 5, we treat age at first birth as exogenous in this estimation, and include the other demographic controls. Identifying instruments include the family background variables as well as nonlabor income and husband's income. Column (8) provides the OLS estimates for the specification and sample used, while column (9) provides the estimates instrumenting for experience and tenure. The Hausman test suggests that there is clear evidence that experience (though not tenure) is endogenous in the wage equation; indeed, the experience coefficient is

[31] We also computed estimates paralleling those in column (7) in which we constructed instruments for the age at first birth dummy variables by estimating an ordered probit for these dummy variables. We used the ranking: age at first birth at age 21 or less; at ages 22–26; at age 27 or higher; and childless. (With age included as a control variable, this seemed the most appropriate ranking for childless.) We used the ordered probit estimates to predict the most likely outcome, and used these expected age-at-first-birth dummies as instruments for the actual first-birth dummies. This procedure (and the procedure described in the text) should lead to consistent estimates and standard errors. However, the estimated standard errors for the age-at-first-birth variables were considerably larger using the ordered probit. The Hausman test also did not reject exogeneity in this specification (p-value = 0.91).

Given the imprecision of the IV estimates, we also experimented with a simpler specification that combined the 27+ and childless as one first-birth category (since their OLS coefficient estimates were quite similar), and dropped the 22–26 dummy. Combining the 27+ and childless categories also allows us to ignore the fertility-timing distinction that might be most directly affected by the current wage error. The OLS coefficient estimate for this group, in a specification corresponding to column (6), was 0.13 (0.03), while the IV estimate (using linear probability instruments) was 0.35 (0.17). However, exogeneity was again not rejected (p-value = 0.18).

(insignificantly) negative in the IV results. These results also suggest larger effects of age at first birth on wages than the corresponding OLS estimates.[32]

V. The relationship between human capital and fertility timing

The results in the previous section indicate that wages are higher for delayed childbearers largely because they have greater accumulation of observable proxies for human capital. As mentioned, this is consistent with our model of joint human capital and fertility timing decisions. In this section, we explore the alternative possibility that the correlation between human capital and fertility timing is spurious, in the sense that human capital and timing appear to be related because both are primarily determined by women's family background.[33]

We would like to be able to disentangle the structural relationship between timing and human capital, but the exclusion restrictions necessary to identify such a model, given our data, would be so arbitrary that an interpretation of the results as valid structural estimates would be highly dubious.[34] Instead, we focus on the "equilibrium" relationship between fertility timing and human capital. In particular, we consider whether the positive relationships between age at first birth and human capital variables in Table 2 are to any extent due to unobserved heterogeneity associated with family background. We estimate regressions of education, experience, and tenure on the same set of age-at-first-birth dummy variables used earlier. We then add an extensive set of family background variables available in the NLS.[35] We do not assert that these variables capture all sources of unobserved heterogeneity. Indeed, if we find that the inclusion of these variables partially reduces the association between fertility timing and human capital, we would have to allow for the possibility that a more complete set of variables could explain the entire relationship. However, if we find no diminution in fertility timing effects once we control for background, it seems reasonable to conclude that heterogeneity does not underlie the results.

The first two columns of Table 6 report results with education as the dependent variable. In column (1) the background variables are excluded, while in column (2) they are included. When these variables are added, the coefficients of the fertility timing dummy variables decline by 20% to 25%. Thus, we cannot decisively reject the view that the education/fertility-timing differentials reflect unobserved heterogeneity rather than human capital investment choices.

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32 Some skepticism might be attached to the IV estimate of a negative return to experience for women. However, it could be argued that the estimates suggest that for women there is no return to experience. Dropping experience as a regressor in the relevant specifications in Tables 3 and 4 had negligible effects on the other estimates, and no effects on the statistical conclusions.

33 Lundberg and Plotnick (1989), McCrate (1989), and Geronimus and Korenman (1992) have considered this possibility for education.

34 One possibility would be to use family background variables as instruments for age at first birth (as at the end of Sect. IV) in regressions with human capital measures as dependent variables. However, we find it particularly difficult to believe that schooling decisions are not directly affected by family background. We are also doubtful that labor-market ties (such as experience and tenure) are not determined partly by family background.

35 Geronimus and Korenman (1992) take this a step further by looking at differences in schooling completion, conditional on whether or not a teen birth occurred, for a sample of siblings. By looking at within-family differences, they may be able to control more thoroughly for differences in family background and other sources of heterogeneity.

Table 6. Years of education, experience, and tenure regressions for white working women, ordinary least squares[a]

	Years of education		Experience		Tenure	
	(1)	(2)	(3)	(4)	(5)	(6)
Age at first birth						
22 – 26	1.59	1.26	0.98	0.97	0.50	0.47
	(0.16)	(0.15)	(0.39)	(0.39)	(0.56)	(0.56)
27 +	3.06	2.30	1.90	1.98	1.01	1.01
	(0.22)	(0.20)	(0.44)	(0.44)	(0.63)	(0.64)
Childless	2.78	2.19	2.00	2.06	1.73	1.74
	(0.15)	(0.14)	(0.37)	(0.37)	(0.53)	(0.53)
Years of education	–	–	−0.69	−0.67	−0.25	−0.19
			(0.07)	(0.07)	(0.09)	(0.10)
Early wage residual[b]	–	–	–	0.39	–	0.84
				(0.32)		(0.46)
R^2	0.27	0.39	0.36	0.37	0.09	0.11
N	1210	1210	701	701	701	701
Family background variables included[c]	No	Yes	No	Yes	No	Yes

[a]Standard errors of estimates are reported in parentheses. Age is included in all regressions.
[b]This is constructed from a wage regression including: years of education and a post-college dummy variable; dummy variables for marital status; dummy variables for residence in the South and in an SMSA; and dummy variables for the year from which the observation was drawn.
[c]Family background variables include: father's education; mother's education; number of siblings; a dummy variable equal to one if the respondent's mother worked when respondent was age 14; a dummy variable equal to one if the respondent lived with both a father and a mother at age 14; and dummy variables corresponding to each of these variables, equal to one when the variable was missing (in which case the variables were set equal to zero).

In columns (3) and (4) we estimate regressions with experience as the dependent variable, and with education included as an independent variable. The equation is identified by assuming that the errors of the education and experience equations are uncorrelated.[36] In the equation for experience in column (4), we also add the early wage residual, to allow for the possibility that delayed childbearers accumulate more experience because they start off with (and possibly continue to have) higher wages. The addition of the early wage residual and the family background variables leaves the estimated age-at-first-birth effects on experience unaltered. In columns (5) and (6) we use tenure rather than experience as the dependent variable; with tenure, inclusion of family background and the

[36] With this restriction, our two-equation model follows the classical recursive-system form. Qualitatively similar results were found using reduced-form experience equations, although the changes in the coefficient estimates when the family background variables are added are more difficult to interpret in this case.

early wage residual reduces the age-at-first-birth coefficient estimates by 10% or less.[37]

We are therefore comfortable concluding that heterogeneity related to family background does not explain the estimated experience and tenure differences associated with fertility timing. Our theoretical model of the relationship between fertility timing and human capital investment offers an explanation for these differences. It may also partially explain the correlation between education and fertility timing, although our results suggest that this empirical relationship is at least partly due to heterogeneity associated with family background characteristics.

V. Summary and conclusions

This paper has developed a model of a woman's optimal human capital investment behavior over the life cycle conditional on her preferences over the timing of her first birth. In the context of this model we show that late childbearers will tend to invest more heavily in human capital than early childbearers. Our model also suggests that women with higher initial wages will choose to delay their childbearing more. Our empirical analysis explores the validity of these theoretical linkages between fertility timing and the wages that women earn while in their late 20's and 30's. Fertility timing is strongly associated with differences in wages, as well as differences in education, experience, and tenure. The wage differences are largely explained by differences in these latter variables, which appear to be good proxies for human capital. We find that the positive relationship between human capital and age at first birth can be only partly attributed to underlying heterogeneity. Thus, the human capital differences seem to explain an important component of the overall relationship between labor market outcomes and fertility timing.

We wish to emphasize that our results are consistent with the human capital hypothesis, but cannot be said decisively to confirm this hypothesis. In fact, what may be a stronger test of our theory – that the usual human capital proxies represent greater investment for delayers – receives little support from the data. Still, our model does provide a unified explanation of the relationships we observe between human capital and fertility timing. We take it as a challenge for future research to develop and test alternative, encompassing models to explain these same empirical relationships.

[37] While we are suspicious of the identifying restrictions, we did estimate experience and tenure equations that excluded the family background variables from the set of regressors, and used them as instruments for the first-birth dummies. The results of Hausman tests suggest that the exogeneity of age at first birth again could not be rejected, with p-values of 0.23 for the experience equation and 0.64 for the tenure equation. The validity of these test results, though, does depend on the appropriateness of the exclusion restrictions.

Appendix A

An age-at-first-birth/wage model with endogenous fertility timing

Notation

Y_I = present value of lifetime income for a woman who invests in human capital
Y_N = present value of lifetime income for a woman who does not invest in human capital
w = wage received at beginning of work career
T = point in time at which woman bears her first (and only) child
τ = length of period spent out of the labor force after childbirth
R = time of retirement
s = growth rate of (real) wages due to investment in human capital
q = growth rate in wages due to general wage growth
r = discount rate

Assumptions

(i) all women are identical (in terms of productivity-related characteristics) at the start of their working career;
(ii) all women work from time 0 to time R, except for the period of length τ following childbirth. τ is the same for all women; in particular it does not depend on T;
(iii) all women have identical discount rates;
(iv) all women have the option of investing in one type of human capital by paying C (at time 0) and receiving a higher growth rate of their wage over time, with the difference in the growth rate of wages between investors and non-investors equaling s;
(v) a woman's lifetime utility depends only on the present value of her lifetime income (i.e., her lifetime earnings) and the age at which she has her first child, i.e., $U = U(Y, T)$, with $\partial U/\partial Y > 0$ and $\partial U/\partial T < 0$. In particular, we assume $U = Y - aT$, with $a > 0$.

A. Fertility timing decisions

(1) Optimal timing for human capital investors. From text Eq. (1), we know that the first order condition for a utility extremum conditional on a woman investing in human capital is:

$$\mathrm{MB}_I(T) \equiv w\mathrm{e}^{(s+q-r)T}[1-\mathrm{e}^{(q-r)\tau}] = a \ . \tag{A.1}$$

Note that if $q < r$, the optimal timing is always $T^* = 0$. For there to be an interior solution for T^*, it is necessary that (A.1) represent the conditions for a maximum and not a minimum. This will be the case if

$$\partial \mathrm{MB}_I(T)/\partial T = w(s+q-r)\mathrm{e}^{(s+q-r)T}[1-\mathrm{e}^{(q-r)\tau}] < 0 \ ,$$

i.e., if $s+q<r$. If $s+q>r$, then either $T^*=0$ or $T^*=R-\tau$.

Assuming $s+q<r$, we have the following description of T^* for investors:

$$T_I^* \begin{cases} = 0 & \text{if } \dfrac{a}{w[1-e^{(q-r)\tau}]} \geq 1 \\[2ex] = \dfrac{1}{(s+q-r)} \log\left[\dfrac{a}{w[1-e^{(q-r)\tau}]}\right] & \text{if } 1 > \dfrac{a}{w[1-e^{(q-r)\tau}]} > e^{(s+q-r)(R-\tau)} \\[2ex] = R-\tau & \text{if } e^{(s+q-r)(R-\tau)} \geq \dfrac{a}{w[1-e^{(q-r)\tau}]} \ . \end{cases}$$

The condition for $T_I^* = 0$ is that $\mathrm{MB}(0) \leq a$; the condition for $T_I^* = R-\tau$ is that $\mathrm{MB}(R-\tau) \geq a$. Note that $\partial T_I^*/\partial a \leq 0$, i.e., the lower the preference for early childbearing the longer the delay before first birth. If $s+q>r$, then the woman chooses to delay until $R-\tau$ as long as $\int_0^{R-\tau} \mathrm{MB}_I(s)\,\mathrm{d}s \geq a(R-\tau)$. This leads to the optimal timing decision:

$$T_I^* \begin{cases} = 0 & \text{if } \dfrac{a}{w[1-e^{(q-r)\tau}]} > \dfrac{e^{(s+q-r)(R-\tau)}-1}{(s+q-r)(R-\tau)} \\[2ex] = R-\tau & \text{otherwise} \ . \end{cases}$$

The upper boundary for T_I^* will hold for lower values of a. [Notice that when $a = w[1-e^{(q-r)\tau}]\dfrac{e^{(s+q-r)(R-\tau)}-1}{(s+q-r)(R-\tau)}$, the woman is indifferent between choosing $T^* = 0$ or $T^* = R-\tau$, since the utility of both choices is equal.]

(2) Optimal timing for non-investors. The first-order condition for a utility extremum [text Eq. (2)] is

$$w e^{(q-r)T}[1-e^{(q-r)\tau}] = a \ . \tag{A.2}$$

The second-order condition for a maximum is satisfied if:

$$w(q-r)e^{(q-r)T}[1-e^{(q-r)\tau}] < 0$$

which holds under the assumption $q<r$. The description of the age at first birth choice is:

$$T_N^* \begin{cases} = 0 & \text{if } \dfrac{a}{w[1-e^{(q-r)\tau}]} \geq 1 \\[2ex] = \dfrac{1}{(q-r)} \log\left[\dfrac{a}{w[1-e^{(q-r)\tau}]}\right] & \text{if } 1 > \dfrac{a}{w[1-e^{(q-r)\tau}]} > e^{(q-r)(R-\tau)} \\[2ex] = R-\tau & \text{if } e^{(q-r)(R-\tau)} \geq \dfrac{a}{w[1-e^{(q-r)\tau}]} \ . \end{cases}$$

As with T_I^*, $\partial T_N^*/\partial a \leq 0$. The condition for $T_N^* = 0$ is the same as for $T_I^* = 0$ (assuming $s+q<r$); however, $T_N^* = R-\tau$ is less likely than $T_I^* = R-\tau$. The expression for the difference in optimal fertility times between investors and non-investors (assuming an interior solution for both) is:

$$T_I^* - T_N^* = \left[\frac{-s}{(s+q-r)(q-r)}\right] \log \left[\frac{a}{w[1-e^{(q-r)\tau}]}\right] > 0 \ ,$$

so that investors will wait longer until their first birth. In addition, $T_I^* = R-\tau$ is more likely than $T_N^* = R-\tau$, which also supports the idea that investors are more likely to delay. It also follows that $\partial(T_I^* - T_N^*)/\partial a < 0$, so that changes in timing preferences have a larger effect on investors' timing decisions than on non-investors'.

B. Human capital investment decision

Given that boundary solutions to the age at first birth decision are possible, it is necessary that we analyze the investment decision under several cases (corresponding to whether or not an investor or non-investor would be at one or the other boundary). There are five separate cases that are exhaustive of the possibilities. (Again, throughout this section we will assume that $q<r$.) Under each case, we derive the following: one, the conditions under which the case holds; and two, an expression for the difference in the indirect utilities $V_j = V_j(w, q, r, a, s, C)$, $j = I, N$, between investors and non-investors. The discussion assumes that the only characteristic that varies across women is the value for a in the utility function.

Case (1): $T_I^* = 0$; $T_N^* = 0$.

Conditions: (A) if $s+q<r$, then this case holds if

$$\frac{a}{w[1-e^{(q-r)\tau}]} \geq 1 \ ;$$

or (B) if $s+q>r$, then this case holds if

$$\frac{a}{w[1-e^{(q-r)\tau}]} > \frac{e^{(s+q-r)(R-\tau)}-1}{(s+q-r)(R-\tau)} \ .$$

Difference in indirect utilities:

$$V_I = \int_{\tau}^{R} w e^{(q-r)t} e^{s(t-\tau)} dt - C$$

$$V_N = \int_{\tau}^{R} w e^{(q-r)t} dt \quad ; \quad \text{so}$$

$$V_I - V_N = \frac{w\mathrm{e}^{-s\tau}}{(s+q-r)}[\mathrm{e}^{(s+q-r)R} - \mathrm{e}^{(s+q-r)\tau}] - \frac{w}{(q-r)}[\mathrm{e}^{(q-r)R} - \mathrm{e}^{(q-r)\tau}] - C\ .$$

The difference in utilities does not depend on a, but lower values of a make this case less likely.

Case (2): $T_I^* = R - \tau$; $T_N^* = 0$.

Conditions: (A) $s+q>r$

$$\text{and (B) } 1 \leq \frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]} \leq \frac{\mathrm{e}^{(s+q-r)(R-\tau)} - 1}{(s+q-r)(R-\tau)}\ .$$

Note that $(\mathrm{e}^x - 1)/x > 1$ as long as $x>0$ (the proof follows from L'Hôpital's Rule and the fact that $\mathrm{d}[(\mathrm{e}^x - 1)/x]/\mathrm{d}x > 0$ when $x>0$), so that this case is possible if $s+q>r$.

Difference in indirect utilities:

$$V_I - V_N = \frac{w}{(s+q-r)}[\mathrm{e}^{(s+q-r)(R-\tau)} - 1] - \frac{w}{(q-r)}[\mathrm{e}^{(q-r)R} - \mathrm{e}^{(q-r)\tau}] - a(R-\tau) - C.$$

It follows that $\partial(V_I - V_N)/\partial a < 0$.

Case (3): $0 < T_I^* < R - \tau$; $0 < T_N^* < R - \tau$

Conditions:

This is the case where neither investors nor non-investors would be at a boundary. This happens if $s+q<r$, and if

$$1 > \frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]} > \mathrm{e}^{(s+q-r)(R-\tau)}\ .$$

Difference in indirect utilities:

$$\begin{aligned} V_I - V_N = {} & \frac{w\mathrm{e}^{-s\tau}}{(s+q-r)}\left[\mathrm{e}^{(s+q-r)R} - \frac{a\mathrm{e}^{(s+q-r)\tau}}{w[1-\mathrm{e}^{(q-r)\tau}]}\right] \\ & - \frac{w}{(q-r)}\left[\mathrm{e}^{(q-r)R} - \frac{a\mathrm{e}^{(q-r)\tau}}{w[1-\mathrm{e}^{(q-r)\tau}]}\right] \\ & + \frac{sw}{(s+q-r)(q-r)}\left[1 - \frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]}\right] \end{aligned}$$

$$+\frac{sa}{(s+q-r)(q-r)}\log\left[\frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]}\right]$$

$$-C\ .$$

The difference in indirect utilities increases as a decreases, since:

$$\frac{\partial(V_I-V_N)}{\partial a}=\frac{s}{(s+q-r)(q-r)}\log\left[\frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]}\right]<0\ .$$

Case (4): $T_I^*=R-\tau;\ 0<T_N^*<R-\tau$

This can happen under one of two sets of conditions:

(A) if $s+q<r$, and $\mathrm{e}^{(q-r)(R-\tau)}<\frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]}\leq\mathrm{e}^{(s+q-r)(R-\tau)}$; or

(B) if $s+q>r$, and $\mathrm{e}^{(q-r)(R-\tau)}<\frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]}<1$.

The difference in indirect utilities is:

$$V_I-V_N=\frac{w}{(s+q-r)}[\mathrm{e}^{(s+q-r)(R-\tau)}-1]-\frac{w}{(q-r)}[\mathrm{e}^{(q-r)R}-1]$$

$$-\frac{a}{(q-r)}\left\{1-\log\left[\frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]}\right]\right\}-a(R-\tau)-C\ .$$

Again, lower values for a will be associated with a greater likelihood of investment in human capital, since:

$$\frac{\partial(V_I-V_N)}{\partial a}=\frac{1}{(q-r)}\log\left[\frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]}\right]-(R-\tau)<0\ .$$

Case (5): $T_I^*=R-\tau;\ T_N^*=R-\tau$

Condition: $\frac{a}{w[1-\mathrm{e}^{(q-r)\tau}]}\leq\mathrm{e}^{(q-r)(R-\tau)}$.

Difference in indirect utilities:

$$V_I-V_N=\frac{w}{(s+q-r)}[\mathrm{e}^{(s+q-r)(R-\tau)}-1]-\frac{w}{(q-r)}[\mathrm{e}^{(q-r)(R-\tau)}-1]-C\ .$$

Conditional on this case holding, the likelihood of investing does not depend on a.

C. Summary

Combining cases where appropriate, it follows that there will exist a single value for $a = a^*$ such that women with $a > a^*$ will choose not to invest in human capital while those women with $a \leq a^*$ will choose to invest. The Appendix table summarizes how the difference in indirect utilities is affected by changes in a. The value for a such that $V_I - V_N = 0$ will be a^*; with lower values of a making higher-numbered cases more likely, it follows that as a is decreasing below a^*, $V_I - V_N$ must be nonnegative and nondecreasing, while as a increases and is greater than a^*, $V_I - V_N$ must be nonpositive and nonincreasing.[38] Lower values of a will thus be associated with more delaying and greater investment in human capital.

Appendix table. *Effect of change in utility parameter a on differences in indirect utility*

Case	$s+q<r$	$s+q>r$
(1)	$\partial(V_I - V_N)/\partial a = 0$	$\partial(V_I - V_N)/\partial a = 0$
(2)	–	$\partial(V_I - V_N)/\partial a < 0$
(3)	$\partial(V_I - V_N)/\partial a < 0$	–
(4)	$\partial(V_I - V_N)/\partial a < 0$	$\partial(V_I - V_N)/\partial a < 0$
(5)	$\partial(V_I - V_N)/\partial a = 0$	$\partial(V_I - V_N)/\partial a = 0$

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[38] For this reasoning to hold, it is also necessary that $V_I - V_N$ not experience discrete downward jumps when moving from one case to another. Continuity of $V_I - V_N$ rules out any such discrete downward jumps. If $s+q<r$, we know $V_I - V_N$ is a continuous function of a, since T_I^* and T_N^* are continuous functions of a, and V_I and V_N are continuous functions of T_I^* and $T_N{}^*$. If $s+q>r$, T_I^* is no longer a continuous function of a; but since the indirect utility of $T^* = 0$ and $T^* = R - \tau$ is the same at the point where the investor switches from case (1) to case (2), it follows that V_I is still a continuous function of a.

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