

**An Economic Model of Teen Motherhood:
Opportunity Costs, Biological Constraints and the
Timing of First Birth**

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ABSTRACT

This paper uses data from the National Longitudinal Survey of Youth to evaluate an economic model of the timing of first birth. We estimate a Tobit-type model, in which a woman chooses her age of first birth based on economic factors subject to the biological constraint that the birth must occur after physical maturity. This model fits the empirical distribution of first birth across a wide range of women. Biological constraints are important, with late maturity associated with low birth rates by age 17 and high birth rates in the later teen years. School achievement, as measured by the Armed Forces Qualification Test (AFQT), accounts for all of the racial differences in timing of first birth and predicts a 10-fold difference in teen-motherhood between top and bottom deciles of achievement. Finally, we find that the model is consistent with self-reports about expected age of first birth and whether a pregnancy was wanted.

I. Introduction

A large literature has documented dramatic differences in fertility patterns by race, education, and income. Some of the most striking evidence involves differences in fertility *timing* across demographic groups and over time, and in particular differences in the timing of first birth. For example, women who are non-white, less well-educated, or living in poverty have first births earlier and are more likely to be teen mothers. Moreover, these differences have been growing over the last three decades, as the well-known increase in the age of first birth – and the related “catch-up” fertility at older ages among recent cohorts – has occurred almost exclusively among white and more educated women (Chen and Morgan, 1991; Rindfuss, Morgan and Offutt, 1996; Morgan et al., 1999; Martin, 2000).

There continues to be much debate over the causes and consequences of these demographic differences in birth timing, and of teen motherhood in particular. On the one hand, an early (e.g. teen) first birth is associated with a wide range of adverse outcomes including dropping out of high school, living in poverty, and being on welfare (Hofferth, 1987). In addition, many teen girls have poor access to or knowledge of effective birth control methods and, correspondingly, teen pregnancies are more likely to end in abortion than pregnancies at any other age (Fielding and Williams, 1991; Ventura, et. al, 1995). These facts have engendered the widely held view that teen motherhood has detrimental effects on a young woman’s future, and is largely the result of mistakes and irrational behavior that ignores these future costs.

On the other hand, recent evidence suggests that the association between teen motherhood and adverse outcomes overstates the causal impact of teen motherhood, because of the self-selection of women with poor labor-market prospects into being teen mothers (Geronimus and Korenman, 1992; Bronars and Grogger, 1992; Hoffman et al, 1993; Olsen and

Farkas, 1989; Ribar, 1994; Rosenzweig and Wolpin, 1995). Moreover, a number of recent micro-econometric studies of the fertility process have found a key role of economic factors, and a woman's wage in particular, in determining the timing of first birth (Heckman and Walker, 1990; Hotz and Miller, 1988). Heckman and Walker in particular found that the strongest effects of economic variables were through the timing of first birth. While some of this evidence has not been focused on teen fertility, it is broadly consistent with an economic model in which a woman chooses an optimal time of first birth, and only chooses to give birth as a teenager when the opportunity costs of teen motherhood are low.

In this paper, we use data on births between the ages of 14 and 37 from the National Longitudinal Survey of Youth (NLSY) to directly investigate the role of economic and other factors in determining the timing of first births. We develop, estimate and evaluate a reduced form economic model in which a woman chooses a target age for first birth based on economic factors (such as benefits of having a child, and opportunity costs in terms of foregone labor market opportunities), and then gives birth at that age subject to biological constraints (e.g. physical maturity). This simple structure implies that the age of first birth can be estimated with a Tobit-type model (with random and unobserved truncation), where both the target age and the truncation point depend on observable variables and unobserved error terms. In our estimation, we allow unobserved heterogeneity in the error terms using a discrete factor approximation (Heckman and Singer, 1984; Goldman, Leibowitz and Buchanan, 1998; Mroz, 1999).

On *a priori* grounds, the model that we estimate has two appealing features. First, economic factors in the model effect a woman's target age of first birth, but her actual age will be later if she reaches physical maturity at a late age. In contrast, a more direct approach relating economic factors to actual age would tend to understate the importance of economic factors

because of the truncation of age of first birth by biological constraints. A second and related feature is that the model's structure directly addresses the issue of whether early births are mistakes (in the sense that they are earlier than the woman wanted). An important feature of mistakes is that if for some reason (such as late maturity) a woman avoids a "mistake" birth, there is no reason for her subsequently to be at higher risk of a birth. In contrast, our model of choice subject to a constraint implies exactly this type of behavior: women whose target age of first birth is below their age of maturity will exhibit catching up behavior once they reach maturity.

We report three primary findings that support our economic model. First, the single strongest predictor of the timing of first birth is a young woman's school achievement (as measured by the Armed Forces Qualification Test). The AFQT is an excellent proxy for opportunity costs (e.g. the foregone uses of the woman's time) in that it is one of the strongest predictors available of both future educational aspirations and labor market earnings. For example, we find that a woman scoring in the bottom decile of the AFQT is ten times more likely to be a teen mother than a woman scoring in the top decile of the AFQT. Moreover, controlling for AFQT reduces or eliminates the effects of other factors, most notably eliminating the effects of race and ethnicity on the timing of first birth. Our second finding is that women who experienced late menarche (i.e. first menstrual period at a late age) were much less likely than other women to have a first birth by age 17, but much more likely to have a first birth between the ages of 18 and 22. Indeed, women who experience late menarche have completely caught up in their childbearing to other women by age 22. This is what one would expect if women were intentionally timing their first birth, subject to the biological constraints imposed by menarche. The final finding is that our simple empirical model, in which women intentionally

choose the time of first birth subject to biological constraints, outperforms a number of more traditional specifications in terms of its ability to reproduce the empirical distribution of first birth across a range of women.

While these empirical results are consistent with an economic model of choice subject to a constraint, they are also consistent with some non-economic models. In the final part of the paper we consider an alternative model, in which the AFQT proxies for poor fertility control and late maturity protects women from the resulting accidental pregnancies. This “accident” model is structurally very similar to the economic model we estimate, and it fits the data on timing of first birth almost as well. However, we find that women with low AFQT scores expect to have an early first birth, are more likely to report that an early pregnancy was wanted, and are less likely to abort an early pregnancy. These results are most consistent with the economic view that teen motherhood among women with low AFQT scores is an intentional choice rather than the result of accidental pregnancies due to poor fertility control.

II. Empirical Model

In this section we develop a simple empirical model of the time of first birth. Our goal is to summarize key features of the distribution, rather than trying to uncover structural parameters. Therefore, we follow Heckman and Walker (1990) in estimating a reasonably parsimonious reduced form model, rather than attempting to estimate a more structural model of fertility dynamics (e.g. Hotz and Miller, 1988). The model we develop differs from more conventional estimation models (e.g. probits models of having a teen birth, proportional hazard models of time to first birth) in that it explicitly incorporates a biological constraint that determines the earliest age at which a woman can give birth.

A. Basic Structure of the Model

A woman is assumed to determine her time of first birth (Y^*) based on two factors: Her target or optimal time for having a first birth (T^*) and her age of reaching reproductive maturity (M^*). Of course there will also be a constraint at the end of reproductive age (menopause), but given the young ages of women in our sample this can safely be ignored. Women whose target age is after they reach reproductive maturity ($T^* > M^*$) will have a birth at their target age. In contrast, women whose target age is before they reach reproductive maturity ($T^* < M^*$) are constrained, and must wait until reaching maturity to have a birth. Assuming some regularity in preferences, the utility loss from delaying birth will be increasing so that these women will give birth immediately on reaching reproductive maturity. Thus, age of first birth will simply be the larger of the target time and reproductive maturity:

$$(1) \quad Y^* = \max(T^*, M^*)$$

Both the target time and the age of maturity are unobserved. Therefore, we model both as a function of observable variables and unobserved, potentially correlated, error terms:

$$(2) \quad T^* = X\beta + \varepsilon^1$$

$$(3) \quad M^* = Z\alpha + \varepsilon^2$$

Since the target age and age of maturity must both be positive, we specify both T^* and M^* as the log of the age.

The structure of equations (1)-(3) is that of a switching model with unknown regimes or, equivalently, a Tobit model with random and unobserved truncation. Identification of this model relies on both exclusion restrictions in the specifications of equations (2)-(3) (e.g. variables that effect target age but not biological maturity, and vice versa) and distributional assumptions on the error terms. We discuss each of these identifying assumptions in turn below.

B. Specification of Equations for Target Age and Age of Maturity

Equation (2) is the reduced-form specification for the target age. The simplest economic models would suggest that either higher benefits from a birth or lower opportunity costs of the woman's time would reduce the optimal time of first birth. Therefore, the determinants of the target age (X) should include proxies for both benefits of a birth (e.g. grew up in a big family, married) and the opportunity cost of a birth (e.g. foregone wages and educational opportunities). Empirically, both wages and education level have been found to be strongly positively associated with age of first birth (Heckman and Walker, 1990; Rindfuss, Morgan and Offutt, 1996). Given our focus on births at young ages, however, variables like these either do not exist in many cases (in the case of current wage) or are very difficult to observe (in the case of likely future educational opportunities). Moreover, these factors and others (e.g. married) are likely to be endogenous to the birth decision (Ribar, 1994; Blackburn, Bloom and Neumark, 1993). Fortunately, in the reduced form, a number of variables are potentially excellent predictors of these opportunity costs. For example, performance on standardized tests such as the AFQT has been shown to be an excellent predictor of earnings later in life as well as eventual educational attainment (Neal and Johnson, 1996). Similarly, education of parents is known to be an important determinant of a child's education (Kane, 1994) as would be predicted from a variety of economic models (e.g. Becker and Lewis, 1973).

Equation (3) is the reduced-form specification for the age of maturity. In principal, the determinants of the age of maturity (Z) should include primarily biological factors that influence a woman's readiness to have a birth. Obviously, one important factor is the age of menarche (when the first menstrual cycle occurred), which ranges widely in the population. However, readiness to give birth is also affected by other physical aspects of puberty that are likely to vary

across women and may occur earlier among some demographic groups. Thus, there is likely to be considerable unobserved heterogeneity in the age of maturity, and some of this heterogeneity may be related to demographic characteristics such as race and ethnicity.

C. Distributional Assumptions

Given distributional assumptions on the error terms (ε^1 , ε^2) in the target age and age of maturity equations, estimation of the model proceeds in a straightforward manner by maximum likelihood. In making distributional assumptions, we address two key issues.

The first key issue in choosing a distribution is to allow for a flexible characterization of unobserved heterogeneity in the target age equation. This is particularly important because the type of catching-up behavior that is implied by our model could potentially be generated by unobserved heterogeneity, e.g. mover-stayer type heterogeneity in which a given fraction of movers rapidly give birth after reaching maturity.

A second key issue in choosing a distribution is to allow for correlation in the errors in the two equations. As documented below, a woman's age of menarche appears to be correlated with various observable measures that we expect to influence target age (e.g. race, family background, and AFQT). Therefore, we might also expect correlation in the unobserved factors influencing these two outcomes. There are a number of plausible reasons to expect a correlation between age of maturity and factors affecting the target age of first birth. For example, age of maturity may affect brain and cognitive development, and thus may influence school performance and learning. Similarly, some evidence suggests that stress during childhood stimulates early physical maturity (Kim and Smith, 1998).

We account for both of these key features by adopting a discrete factor approximation for the error terms. This follows the non-parametric approach of Heckman and Singer (1984) in

assuming that unobserved heterogeneity (θ) takes on a discrete number of values, with the number of points of support (K), the values ($\theta_1, \dots, \theta_k$) and associated probabilities (π_1, \dots, π_k) all estimated empirically. Moreover, any correlation between the two errors ($\varepsilon^1, \varepsilon^2$) is assumed to depend on this unobserved factor through:

$$(4) \quad \varepsilon^1 = \theta + v^1, \quad \varepsilon^2 = \rho\theta + v^2, \quad \text{where } (v^1, v^2) \sim N\left(0, \begin{bmatrix} \sigma_1^2 & 0 \\ 0 & \sigma_2^2 \end{bmatrix}\right)$$

Thus, conditional on the unobserved factor (θ), target age and age of maturity are distributed independently and follow a log normal distribution. The unobserved factor both allows for deviations from the log normal distribution, and allows for correlation across the two distributions through ρ . This approach has been shown to work successfully in monte carlo simulations (Mroz, 1999) and in replicating experimental estimates with observational data (Goldman, Leibowitz and Buchanan, 1998).

D. Estimation.

Estimation of the model is by maximum likelihood, where the likelihood must further account for censoring of the time of first birth for women who had not yet had a birth by the end of the data period. Let Y equal the censored value of (log) age at first birth, Y^0 be the age of censoring, and let D be a dummy variable indicating if the data were censored, i.e.:

$$(5) \quad Y = \min(Y^*, Y^0), \quad D=1(Y^* > Y^0)$$

The CDF of Y^* , conditional on the unobserved factor (θ), is simply:

$$(6) \quad \begin{aligned} \Pr(Y^* < y \mid \theta) &= \Pr(\max(T^*, M^*) < y \mid \theta) \\ &= \Pr(T^* < y \text{ and } M^* < y \mid \theta) \\ &= \Pr(T^* < y \mid \theta) \Pr(M^* < y \mid \theta) = \Phi\left(\frac{y - X\beta - \theta}{\sigma_1}\right) \Phi\left(\frac{y - Z\alpha - \rho\theta}{\sigma_2}\right) \end{aligned}$$

Where the last step follows from the fact that the errors are independent normal conditional on θ .

The contribution to the likelihood of the censored variable (Y) conditional on θ can be derived in a straightforward manner as:

$$(7) \quad L|\theta = D \left[1 - \Phi \left(\frac{y - X\beta - \theta}{\sigma_1} \right) \Phi \left(\frac{y - Z\alpha - \rho\theta}{\sigma_2} \right) \right] \\ + (1 - D) \left[\frac{1}{\sigma_1} \phi \left(\frac{y - X\beta - \theta}{\sigma_1} \right) \Phi \left(\frac{y - Z\alpha - \rho\theta}{\sigma_2} \right) + \frac{1}{\sigma_2} \Phi \left(\frac{y - X\beta - \theta}{\sigma_1} \right) \phi \left(\frac{y - Z\alpha - \rho\theta}{\sigma_2} \right) \right]$$

Integrating across values of θ yields the unconditional likelihood:

$$(8) \quad L = \sum_{k=1}^K \pi_k \left\{ D \left[1 - \Phi \left(\frac{y - X\beta - \theta_k}{\sigma_1} \right) \Phi \left(\frac{y - Z\alpha - \rho\theta_k}{\sigma_2} \right) \right] \right. \\ \left. + (1 - D) \left[\frac{1}{\sigma_1} \phi \left(\frac{y - X\beta - \theta_k}{\sigma_1} \right) \Phi \left(\frac{y - Z\alpha - \rho\theta_k}{\sigma_2} \right) + \frac{1}{\sigma_2} \Phi \left(\frac{y - X\beta - \theta_k}{\sigma_1} \right) \phi \left(\frac{y - Z\alpha - \rho\theta_k}{\sigma_2} \right) \right] \right\}$$

Estimation of the parameters is by maximum likelihood, with the number of points of support being chosen based on improvements in the likelihood function.

Finally, we parameterize the model in a way that imposes some necessary normalizations. The standard deviations (σ_1, σ_2) are estimated as logs to ensure that they are positive. To ensure that the probabilities sum to one over the points of support, we estimate a multinomial logit form, where:

$$(9) \quad \pi_k = \frac{e^{\delta_k}}{\left(1 + \sum_{j=1}^{K-1} e^{\delta_j} \right)} \text{ for } k = 1 \text{ to } K-1.$$

Lastly, we include an intercept in $X\beta$ and normalize $\theta_1=0$.

III. Data

The data comes from the National Longitudinal Survey of Youth (NLSY). The NLSY is a panel survey of roughly 12,000 individuals (men and women) who were first surveyed in 1979 when they were between the ages of 14 and 21. Our sample consists of 4224 women who

remained in the survey through 1994 (when they were ages 29-37) and had complete data on all the key variables. The NLSY over samples minority and poverty populations. The estimates reported in this paper have not been weighted by sampling weights, and therefore may not be representative of the U.S. population. For details of the data construction, see the attached data appendix.

Information on the age of first birth is derived from detailed birth history information asked every year of the survey. Any birth that occurred prior to 1994 (including those that occurred prior to the start of the survey in 1979) is in principal recorded. Questions about age of menarche were asked in 1984 and 1985, when the sample was at least 18 years of age. A few women had not yet reached menarche by this time, and their age of menarche was coded as 18.

The NLSY has also asked questions about every pregnancy, beginning retrospectively in 1982. Each woman was asked for the date at which each pregnancy began and the outcome of each pregnancy (abortion, live birth, miscarriage, stillborn). We use these questions to determine the age of first pregnancy and whether the pregnancy ended in abortion. Additional questions about each pregnancy were asked, including: was the woman using contraception at the time of becoming pregnant? If not, was that because the pregnancy wanted? And finally all women who had not yet indicated that they wanted to become pregnant were asked a follow-up question, “Just before you became pregnant, did you want to become pregnant?” We use these questions to designate each woman’s first pregnancy as wanted if she indicated that the pregnancy was “wanted” or “didn’t matter” on either the initial or the follow-up question.

Information on achievement comes from the Armed Forces Qualification Test (AFQT), an achievement test that was administered to each woman in our sample in 1980. The AFQT is based on a subset of tests from the Armed Services Vocational Aptitude Battery (ASVAB) that

have been developed by the military for enlistment screening and job placement. The AFQT has been validated in a number of studies, and found to be both racially unbiased and an excellent predictor of future job performance, education and labor market outcomes (see Neal and Johnson, 1996, for a careful summary of the evidence). We use the more recent (1989) scoring of the AFQT, and use the percentile score (based on national percentiles) as our measure. In all specifications reported in this paper, we have found effects of the percentile score to be remarkably linear. Finally, the AFQT was administered to some women in our sample after the time of their first birth. This is potentially problematic, since having an early first birth is likely to affect subsequent skill acquisition and therefore also affect AFQT performance. Therefore, in some analyses we limit the sample to women who were age 14 or 15 in 1979, since very few of these women had births prior to taking the AFQT in 1980.

Other variables are derived primarily from questions asked in 1979. We use a question about expected time until first birth to calculate expected age of first birth. This variable is missing for women who had already had a birth by the 1979 interview, or who expected to have no children. We use self reported race and ethnicity to categorize the sample into black, non-black Hispanic (which we refer to as Hispanic), and all others (which we refer to as White). Questions about mothers and fathers education were used create dummy variables for whether each individual's mother (father) graduated from high school or college. Individuals who did not respond because their mother/father was not present were included in the base category (did not graduate from high school). Finally, we use questions from 1979 to determine: the number of siblings in the family; whether the individual was living in a single parent family at age 14; and whether the individual had no reading material (magazines, newspapers, or library card) in the house at age 14, or had all of these (with the base case being having some).

IV. Results

We divide our results into our results into five parts. We begin with some descriptive evidence establishing the basic facts in the data. The second part uses reduced form probit models to investigate the importance of economic factors in determining teen motherhood. The third part uses similar probit models to investigate the extent that women with late menarche have increased rates of first birth in subsequent years. The fourth part of our results discusses estimates of the empirical model developed in section II, and evaluates the ability of this model to fit the data relative to other more conventional models of age at first birth. Finally, we evaluate the ability of an alternative non-economic model to fit the data on timing of first birth, and investigate which model is more consistent with a woman's statements about her expected age of first birth, whether her first pregnancy was wanted, and whether her first pregnancy ended in abortion.

A. Descriptive Results

Table 1 provides summary statistics on variables used in the analysis for the sample as a whole, and separately for women who experienced menarche at an early (11 or younger), average (12-14), or late (15 or older) age. Perhaps the most striking fact in the table for those who are not familiar with this data is the rate of early childbearing: over 13% of the sample has had a birth before age 18, and over a quarter before age 20, with the median first birth occurring at age 24. These rates are a bit higher than national averages, reflecting the over sampling of minority and poverty populations in the NLSY.

Given the early childbearing in this population, biological maturity is potentially a meaningful constraint on childbearing. The average age of menarche in the sample is just under

13 years, but roughly 15% of the sample either has menarche before age 12 or after age 14. Timing of menarche appears to have a strong effect on birth timing at young ages: Women with early menarche are nearly twice as likely as women with late menarche to have a first birth before age 18 (15.8% versus 7.4%). However, nearly all of this difference has disappeared when looking at the percent of these women having a birth before age 23. This is precisely the kind of catching up behavior that is consistent with menarche acting as a constraint on early childbearing that is intentional (as opposed to accidental).

However, there are at least two reasons that this simple evidence may not be conclusive. First, this type of catching up can be the result of heterogeneity in the population. For example, a standard hazard model with mover-stayer heterogeneity could generate this pattern. In addition, age of menarche does not appear to be randomly distributed. As seen in Table 1, women experiencing late menarche have lower AFQT scores, more siblings, are more likely to be black, less likely to be Hispanic, and more likely to have lived in a single parent household at age 14. Note also that the relationship is not always monotonic, with early menarche women also having lower AFQT scores and more likely to be black. At the very least, this suggests that some of the relationship between menarche and timing of first birth may be due to unobserved differences between these women.

While menarche appears to be related to early fertility, the effects are small relative to other factors that are associated with timing of first birth. Figure 1 displays Kaplan-Meier estimates of survival curves (the proportion of women not having a first birth at each age) for our sample, broken down by race and ethnicity, sample decile of the AFQT (not national deciles), and by age of menarche (after age 14 versus all others). The survival curves by late menarche clearly display the catching-up phenomenon, briefly diverging in the teens before coming back

together just after age 20. However, differences by age of menarche are small relative to differences along other dimensions. As has been well documented elsewhere, blacks and Hispanics are more likely to have had a birth at early ages (although these differences shrink at older ages). More striking, however, are the differences in timing of first birth across deciles of the AFQT: The difference between the top and bottom decile in probability of a first birth before age 20 is more than 10-fold (under 5% versus over 50%), with the survival curves moving uniformly downward at younger ages as one moves down in the AFQT distribution. This evidence is consistent with that of Heckman and Walker (1990) and others suggesting that women with high returns in the labor market tend to delay first birth.

B. Effects of Economic Factors on Early Childbearing

Table 2 investigates the factors associated with early first birth in more detail. Table 2 estimates probit models (with probability derivatives given in square brackets) of the determinants of teen birth (age 19 or younger). We estimate models for the entire sample and in a sample limited to women who were age 14 or 15 in 1979. The latter estimates address the concern that performance on the AFQT may be affected for women who had an early first birth prior to taking the test. For each sample, we estimate three models. The first specification only includes black and Hispanic dummies, the second adds the AFQT percentile, and the third adds a variety of family background measures.

Four important conclusions can be derived from these estimates.

The first conclusion is that the AFQT score is the single strongest predictor of early childbearing, as measured by either statistical significance or magnitude of the effects. In models using the entire sample and controlling for family background, we estimate that a one-percentile increase in the AFQT is associated with 0.4 percentage point reduction in teen

pregnancy (column 3). These effects are even larger in specifications that do not control for family background.

A second conclusion is that differences in the timing of first birth for blacks and Hispanics are largely eliminated in models that control for AFQT. For example, being black is associated with a 23.4 percentage point higher rate of teen birth when we do not control for AFQT, but this estimate falls to 7.8 percentage points when we control for AFQT and even further to 3.6 percentage points when we control for family background. As one would expect if wages were a key determinant of age at first birth, these results are consistent with Neal and Johnson (1996), who find that *wage* differences between white, black and Hispanic women are eliminated once one holds performance on the AFQT constant. Thus, as found by Neal and Johnson, it appears that academic achievement (or skill) differences can explain other large racial and ethnic differences.

A third conclusion derived from Table 2 is that the potential bias from women taking the AFQT at older ages (after first birth) is apparently small in practice. Estimates for the full sample are very similar to estimates based on a sample of women who were 14 or 15 in 1979. In fact, if anything, the effects of AFQT are larger in the restricted sample, which is the opposite of what one would expect if early childbearing led to lower subsequent AFQT scores. Thus, for the remainder of the paper we focus on estimates for the full sample.

C. Effects of Constraints on Early Childbearing

The descriptive evidence in Table 1 and Figure 1 suggest that women experiencing late menarche (1) are less likely to have a first birth at very young ages, but (2) are more likely to

have a first birth subsequently and are largely caught up by their early 20s. Table 3 investigates this fact in more detail.

In Table 3 we report probit estimates of models that include a dummy variable for whether the woman had a late menarche (after age 14), and that control for race, AFQT and family background. The first two columns are estimated on the full sample. The first column reports estimates in which the dependent variable is whether the woman had a first birth before age 18. In the second column the dependent variable is whether the woman had a first birth between the ages of 18 and 22. Note that the second column includes all women, so is estimating the unconditional probability (not the hazard rate). Thus, these models provide estimates of the extent to which late menarche reduces the probability of early first birth (<18) or increases the probability of first birth at subsequent ages (18-22). The remaining columns of the table repeat these specifications for women in the bottom quartile (nationally) of the AFQT and women above the bottom quartile.

The magnitude of the effect of late menarche in the full sample is very similar to what was observed in the means reported in Table 1. Late menarche is associated with a 6 percentage point reduction first birth before age 18, and a corresponding 6 percentage point increase in the probability of first birth between 18 and 22. Both estimates are highly significant. Thus, women with late menarche are completely caught up by age 22 in the cumulative probability of first birth.

If late menarche is acting as a constraint on early childbearing, then we should observe the most pronounced effect of late menarche on women who are most likely to have a birth in their early teens. Therefore, in the remaining columns of Table 3 we split the sample according to whether women scored in the bottom quartile of the AFQT. As expected, the impact of late

menarche is seen primarily among women with low AFQT scores. For these women, late menarche is associated with a 12 percentage point reduction in first birth before age 18, and a corresponding 15 percentage point increase in the probability of first birth between 18 and 22. For women who are not in the bottom quartile of the AFQT, we estimate a smaller (in absolute but not relative terms) and significant effect of late menarche on having a first birth before age 18, but then no catching up. The effect of late menarche is actually negative (but insignificant) for the high AFQT women on the probability of first birth between 18 and 22. One possible interpretation of these findings is that early births among high AFQT women are primarily mistakes (with no associated catch-up), while early births among low AFQT women are intentional (with the associated catch-up).

D. Estimation and Evaluation of Our Empirical Model

The preceding sections lend general support to the key idea embedded in the empirical model developed in section II – namely that the timing of first birth can be characterized by women choosing a target age that depends on economic factors, subject to biological constraints. In this section we estimate the model, and evaluate the model’s ability to fit the data relative to two more conventional models: a Cox proportional hazard model, and a log normal that incorporates heterogeneity using the discrete factor approach (but which does not explicitly incorporate the biological constraint).

Table 4 contains maximum likelihood estimates for a variety of models and specifications. In addition to variables reported in the table, all models control for the full set of family background measures listed in Table 1. While we do not report them here, parents’ education had the largest and most significant effects on age of first birth.

The first column of Table 4 reports coefficient estimates from a Cox proportional hazard model of the time to first birth, where the hazard rate is proportional to $\exp(X\beta)$. The next two columns report estimates for a standard log-normal model, in which the log of the age of first birth follows a normal distribution, with mean $X\beta$ and standard deviation σ . This is a special case of our model, in which the constraint is assumed to be non-binding (e.g. the mean of the biological constraint is at a very low age). The second column is estimated with no heterogeneity, and the third column allows for heterogeneity with four points of support. Adding additional points of support did not improve the likelihood or change the relevant estimates. The remaining columns of Table 4 report estimates of the constrained log normal model that was developed in section II. Columns 4 and 6 estimate the model with no unobserved heterogeneity, while columns 5 and 7 estimate the model with discrete factor heterogeneity with two points of support. Once again, additional points of support did not improve the likelihood or materially affect the reported results. Note that the coefficients from the Cox model are not comparable to coefficients from the remaining models: a variable that increases the cox hazard rate will reduce the mean age of first birth in the remaining models.

Estimates from the Cox proportional hazard model in the first column of Table 4 are similar in character to the estimates from previous tables. The AFQT is a strong predictor of delayed childbearing (reducing the hazard rate). The coefficients for black and Hispanic are both negative but only marginally significant (lowering the hazard, in contrast to the much higher birth hazard observed for black and Hispanic when not controlling for AFQT). In this model, we have added age of menarche (in log form) directly as a control variable. Age of menarche is estimated to have a small but significant effect, with late menarche associated with a lower birth hazard.

The proportional hazards assumption, however, is strongly rejected by the data. Tests of the proportional hazard assumption (Grambsch and Therneau, 1994) reject at the .001 confidence level. The reason for this rejection is apparent in Figure 2a, which plots the empirical distribution (derived from Kaplan-Meier estimates of the survival function) of age of first birth for each sample decile of the AFQT, along with the distribution predicted by the Cox model. The predicted distributions from the Cox model is derived for each person based their individual characteristics and the parameter estimates, and then these distributions are averaged over the entire sample to create the estimates in Figure 2. The most notable feature of the empirical distribution is that the mode is clearly shifting to the right as AFQT increases. Proportional hazard models, which require the AFQT and other variables to shift the hazard up or down uniformly at all ages, are not able to capture this feature of the data.

Estimates from the log-normal model in the second column of Table 4 are similar in character to the estimates from Cox hazard model. The coefficient on the AFQT implies that for every 10 percentile (e.g. 1 decile) increase in the AFQT, there is an approximately 2.8% increase in the median age of first birth. Based on the median age of first birth in the sample (24 years), this translates into an increase of roughly 2/3 of a year for every decile increase in the AFQT. Estimates of the effect of black and Hispanic are both positive, and the black effect is insignificant. Age of menarche is estimated to have a small but significant effect, with a 10 percent increase in age of menarche estimated to increase the median age of first birth by just under 1 percent.

Results in the third column of Table 4 are fairly similar. The most notable changes are a slightly negative and significant coefficient on black, and a larger effect of menarche. This specification has added discrete heterogeneity with four points of support (six new parameters),

and the resulting gain in the likelihood is very significant. While we do not report these parameters, they are generally precisely estimated and imply points of support around 20, 24, 33 and near 80 (imprecise, but essentially implying never). Most of the probability mass is on the first two points (about 40% each), with the remainder split evenly on the two higher points.

The remaining columns of Table 4 report estimates from models that incorporate the biological constraint. In all specification, we let the biological constraint depend on age of menarche, and black and Hispanic dummy variables. In columns 4 and 5 we do not include menarche in the target age equation, while in the last two columns we put menarche in the target age equation as a check on the assumptions of the model.

There are a number of notable differences between the standard log normal model and the constrained log normal model. First, the likelihood function improves significantly relative to the few additional parameters being estimated. For example, in models without discrete factor heterogeneity (column 2 versus column 4 or 6) the likelihood in the constrained log-normal model improves by more than 300. For models that incorporate heterogeneity, the improvement is not as dramatic, but still the constrained log-normal model clearly dominates the standard log normal model.

Another difference in the constrained log-normal estimates is that the estimated impact of AFQT on target age increases. Depending on specification, these estimates imply that a one decile increase in the AFQT is associated with an increase in the median age of first birth of between .75 and 1 year. In other words, the censoring of target age in the standard log-normal model leads to an underestimate of the full effect of AFQT on target age. Similarly, the estimated standard deviation around the median target age is smaller in the standard log-normal model because of the censoring. As a result of the larger estimated effects of the AFQT, the

black and Hispanic effects have become more positive, but remain generally small and often insignificant.

Estimates of parameters in the constraint equation are also interesting. Age of menarche is positive and significant as expected, with a 10% increase in age of menarche associated with an increase in the constraint age of between 3% and 4%. Moreover, when menarche is added to the target age equation, its effect is relatively small and insignificant. Thus, the model is able to accurately identify menarche as a variable that works through the constraint. Being black also appears to work through the biological constraint, reducing the biological constraint by roughly 1 year. This estimate is consistent with some evidence suggesting that blacks have earlier physical development. Note that one of the reasons that race has less of a negative effect on target age in this model, as compared to the standard log-normal model, is that this model is estimating that race works through biological constraints rather than through target age of first birth (holding AFQT constant).

Finally, the constrained log-normal models that incorporate discrete factor heterogeneity provide fairly similar estimates to those without heterogeneity. As in the standard log-normal model, introducing heterogeneity reduces the effect of AFQT slightly, makes black effects less positive, and increases the effect of menarche. Introducing the discrete factor heterogeneity (which only adds three parameters) does improve the likelihood significantly. As in the standard log-normal model, we estimate that about 80% of the sample is at a point with median target age of first birth of around 21, while the remaining 20% have a median target age around 44. Moreover, we estimate that this heterogeneity has no significant relationship to the error in the constraint equation.

Overall, these estimates are consistent with the original motivation of our empirical model. We find that a woman's performance on the AFQT (and to a lesser extent parental education) is the single most important predictor of her target age of first birth. Both factors are important determinants of educational and labor market opportunities. Moreover, by incorporating age of menarche as a constraint, rather than a factor shifting optimal time, we significantly improve the fit of the model.

The constrained log-normal model is also able to predict important features of the distribution of first birth that are not predicted by other models. Figure 2b plots the empirical distribution of first birth in the ten sample AFQT deciles along with the predicted distribution derived from the constrained log-normal model with discrete factor heterogeneity. The constrained log-normal model replicates the empirical distribution remarkably well. In particular, the model can replicate both the large spike at young ages for low AFQT deciles as well as the less pronounced mode at higher ages for high AFQT deciles. In contrast, the standard log-normal model with discrete heterogeneity (but without the constraint) cannot fit the data as well, as can be seen in Figure 2c. As with the Cox model, the standard log-normal model is unable to predict the large spike at low ages for the low AFQT deciles.

To evaluate the predictive ability of the models more quantitatively, Table 5 runs simple regressions in which the dependent variable is the empirical probability of having a first birth in a cell defined by single year of age and AFQT decile (e.g. the same points as are plotted in Figure 2). The right hand side variables are the predicted probability from each model for each of these cells. For this table, we focus only on the better fitting models with discrete heterogeneity. We report OLS estimates, but WLS estimates were very similar. The constrained log-normal model is best able to predict the empirical probability of having a first birth in each of these cells.

When included separately the constrained log-normal model yields a higher R-squared and has a coefficient not significantly different from 1. Moreover, when we add predictions from other models, the constrained log-normal predictions are always the dominant predictor (although the Cox model has some additional predictive power).

Table 6 provides an alternative measure of each model's ability to fit the data. For each model, this table reports chi-squared goodness of fit statistics in each decile of the AFQT and overall. These measures were used by Heckman and Walker (1990) to evaluate fertility models in a similar context. The test statistic is given by:

$$R = N \sum_{j=1}^C \frac{(\hat{p}_j - p_j)^2}{\hat{p}_j}$$

where N is the overall sample size across all cells, C is the number of cells, \hat{p} is the model's predicted proportion of the sample in that cell, and p is the actual proportion of the sample in that cell. If the model is the true model generating the data, this statistic has a chi-squared distribution with C-1 degrees of freedom, and lower values indicate a better fitting model. To avoid small cells, we define age cells from 16 to 29, with both the end points including births at earlier/older ages. Thus for each decile of the AFQT, we are fitting 14 cells and have 13 degrees of freedom. Since the statistics are independent across AFQT deciles under the null, we can add the statistics across deciles to form an overall test statistic with 130 degrees of freedom.

As seen in Table 6, the constrained log-normal model provides the best overall fit of the data, although the overall goodness of fit statistic rejects all three models. Looking at the results for individual deciles of the AFQT, the constrained log-normal model outperforms the standard log normal model in every decile, and outperforms the Cox model in 9 of 10 deciles. Moreover, the constrained log-normal model passes the goodness of fit test in eight deciles (at the 5% level)

and is close in one more. In contrast, the Cox model only passes in three deciles and is close in two more, while the standard log-normal model only passes in four deciles and is close in one more. In general these tests confirm what was seen in Figures 2a-2c: The constrained log-normal model is the only model that is able to fit the data well throughout the AFQT distribution.

A final comparison of the models ability to fit the observed data is provided in Figures 3a-3c. These figures compare each model's predictions to the actual distributions of first birth when the sample is split into four sub-samples according to whether the woman was in the bottom quartile of the AFQT, and whether the woman had menarche after age 14. Thus, these figures ask which of the models are able to replicate the catching-up patterns of first birth documented in Table 4. Once again, it is apparent that only the constrained log-normal model is able to fully capture the patterns observed in the women with late menarche. In particular, the constrained model is able to almost perfectly predict the low initial probability of a first birth followed by a rapid spike in the probability for women with low AFQT and late menarche.

D. Evaluation of an Alternative Model of Accidental Pregnancy

Given the empirical success of our simple economic model, it is natural to ask whether there are other models that provide an equally good fit of the data. In particular, is there a non-economic model that can explain both the importance of the AFQT and the catching up in birth rates observed among late menarche women? In this section we consider one such model, in which the AFQT proxies for poor fertility control and late maturity protects women from the resulting accidental pregnancies. This "accident" model is structurally very similar to the economic model we estimate, and it fits the data on timing of first birth almost as well.

However, we argue that the economic model is more consistent with other data on early pregnancies.

Suppose that teen births are largely the result of accidental pregnancies, for example resulting from a limited understanding of contraception, or poor communication of the woman's preferences to her partner (e.g. because of low self-esteem), or an inability to act on those preferences (e.g. because she feel required to defer to her partner's wishes, or is simply unable to sort out what she takes to be conflicting ideals or demands). In this case, one might expect that women with low school achievement (or from disadvantaged backgrounds) would be most at risk for accidental teen pregnancies. Moreover, late menarche would protect such women against the risk of accidental pregnancies. However, after reaching maturity, these women would be at risk for accidental pregnancies at their usual high rate, so that late menarche would only provide temporary protection.

The "accidents" model is structurally very similar to our economic model, and can be estimated using a slightly modified likelihood function. Both models imply that low AFQT will be associated with early birth, subject to a biological constraint. The only difference is that the accident model assumes that the birth *hazard* for an individual woman will be unaffected by late maturity. Thus, rather than assuming that the distribution of age of first birth conditional on age of maturity is a truncated normal (which includes a spike at the age of maturity), we assume that the distribution is a censored normal (which eliminates the spike, and scales up the remainder of the distribution proportionally). Although this accidents model has no catching up in the birth rate of an individual following late maturity, it can generate catching up in aggregate if there is unobserved heterogeneity in the hazard rate across individuals.

Parameter estimates of this accident model look quite similar to those of the constrained log-normal model, and the accident model does only slightly worse in terms of fitting the data. The log likelihood for the accident model with discrete factor heterogeneity and two points of support is -971 , only slightly worse than the constrained log-normal model's log likelihood of -969 , and the Vuong non-nested likelihood test cannot reject either model in favor of the other. The accident model also does slightly worse than the economic model in fitting the empirical distributions of age of first birth by AFQT decile (as done in Table 6). The overall goodness of fit statistic for the accident model is 187.81 (versus 178.58 for the economic model), and fit of the accident model is worse in seven of the AFQT deciles, including the five lowest deciles.

The similarity of the two models is clearly apparent in Figure 4, which compares each model's predictions to the actual distributions of first birth by AFQT and late menarche, as was done in Figures 3a-3c. Only among the women with low AFQT and late menarche does the economic model fit the data noticeably better, capturing slightly more of the spike in births for these women in their late teens. Thus, as with the comparisons of likelihood and goodness of fit statistics, the evidence favoring the economic model is not overly compelling.

Although the two models appear to have very similar implications concerning the timing of first birth, they have quite different implications in terms of whether teen pregnancies and resulting births are planned and wanted. In the economic model, women with low AFQT plan to have a teen birth, get pregnant as teens intentionally, and a large proportion of the resulting pregnancies, therefore, will be wanted and carried to term. In contrast, the accident model implies that women with low AFQT do not plan to have a teen birth, get pregnant accidentally, and a large proportion of these pregnancies are unwanted and will be aborted. In the remainder of this section we investigate which model is more consistent with a woman's statements about

her expected age of first birth, whether her first pregnancy was wanted, and whether her first pregnancy ended in abortion.

In Table 7 we estimate the relationship between the target age estimated from the economic model (the constrained log-normal model) and the expected age of first birth that women reported in 1979. There are three limitations of the data on expected age that limit this analysis. First, the question asks about “expected” age of first birth, which may not necessarily be the age at which a woman wants to have her first birth. Second, there is no data for any woman who had already had a child by 1979. Therefore, we estimate the models on both the full sample and on the sample of women who were 14 or 15 in 1979. Less than 1% of the latter sample had a child by 1979. Finally, roughly 10% of the sample did not answer the question either because they didn’t know or because they expected to have no children. We have left these women out of the analysis, and made no attempt to correct for the resulting sample selection bias.

The results in Table 7 suggest that women with low AFQT expect to have an early birth, and that the target age from the economic model is a good predictor of these expectations. The first and third columns of Table 7 report the result from a simple regression of the log of the expected age of first birth on the log of the target age of first birth ($X\beta$) as estimated from the economic model. The first column uses only the sample of women who were 14 or 15 in 1979, while the third column uses all women. In both samples, there is a positive and significant relationship between the target age and what women report as their expected age of first birth, although the coefficient is well below its expected value of 1. The second and fourth column run an unconstrained regression, allowing all of the X ’s to enter individually. In neither sample does this add much to the explanatory power of the regression as measured by the R-squared,

although the additional variables are significant in the full sample. In other words, the target age serves as a summary statistic in predicting a woman's expected age of first birth. In these unconstrained models, the AFQT is positive and significant, with an increase of one decile in the AFQT associated with a 1% (or about one quarter year) increase in expected age of first birth.

Finally, we investigate the relationship between target age estimates derived from our model and the likelihood that a pregnancy is wanted. An implication of the economic model is that a woman should be more likely to indicate that she did not want the pregnancy if we estimate that the pregnancy occurred well before her target age. Conversely, women who have pregnancies at young ages, but for whom we estimate a young target age, should be more likely to indicate that the pregnancy was wanted.

To investigate this question, we relate a woman's report of whether her first pregnancy was wanted to our model's prediction of the log difference between her age at the time of the pregnancy (plus 1 to give age at birth) and her target age. When this difference is negative, pregnancy is likely to have occurred too early and will therefore be unwanted. We focus on first pregnancies occurring by the 1984 interview in order to focus on pregnancies at a young age. The first column of Table 8 reports the number of pregnancies in our sample at each age. The second column reports the proportion of women at each age that reported their pregnancy was wanted. The final column reports the proportion of pregnancies at each age that were not aborted, which serves as a more concrete indication of whether the pregnancy was wanted. A large proportion of the pregnancies are wanted, ranging from around one quarter of the pregnancies under age 14 to over 70% by the mid 20s. Very few pregnancies are reported to end in abortion at any age, which most likely is due to under-reporting of abortions in the NLSY (Jones and Forrest, 1992).

Table 9 reports regressions of these two measures of pregnancy wantedness on our estimate of whether a pregnancy was after the target date and other factors. The first column includes only the log difference between pregnancy age and target age. The coefficient on this variable is positive and strongly significant, as expected, implying that a 10% increase in the age of pregnancy (e.g. from 18 to 20) is associated with a 6.7 percentage point increase in wantedness. Adding race dummies in column 2 does not have much impact on this estimate, although the black dummy is highly significant and suggests that black women are much less likely to report a pregnancy is wanted. In column 3 we run the model unrestricted, including all of the X's that determine the target age plus the log age of the pregnancy. This improves the R-squared only slightly, although one can formally reject the restrictions imposed by using the target age variable. More importantly, AFQT remains a strong negative predictor that the pregnancy is wanted in the unconstrained model. The results in the remaining three columns of Table 9 repeat this analysis for whether the pregnancy was not aborted, and the findings are quite similar.

Overall, the evidence based on a woman's stated preferences about pregnancies and related actions regarding abortion favors an economic model, in which women intentionally choose to give birth at young ages, over a model in which teen births are accidental. In particular, the women that the model predicts as likely to have teen births report expectations of an early birth, are more likely to report that a teen pregnancy was wanted, and are less likely to abort a teen pregnancy.

V. Conclusion

The evidence we present in this paper provides further support for the economic model of fertility. In particular, we believe that the distribution of first birth at early ages has features that are difficult to explain with conventional models, but are natural implications from a model in which a woman chooses her optimal time for first birth subject to constraints. Based on this model, we find that the key predictor of optimal age of first birth is a woman's achievement or skill as measured by the AFQT. Race is found to have little or no direct effect (controlling for AFQT) on optimal age, but does have some direct effect on timing of first birth through racial differences in the biological constraint.

These results lend further support to the argument that teen motherhood is not generally a mistake, but rather intentional behavior on the part of young women who face low opportunity costs of a birth. While efforts to improve information available to teens on birth control and the consequences of a teen birth are, no doubt, effective to some degree, our evidence suggests that efforts to improve school achievement and labor market opportunities are the key to lowering rates of teen motherhood.

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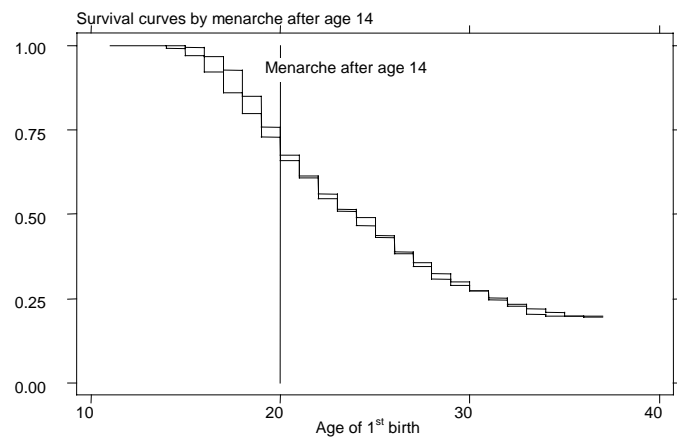
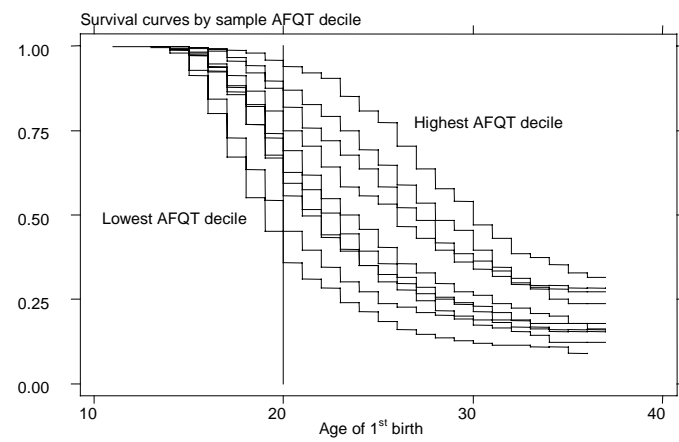
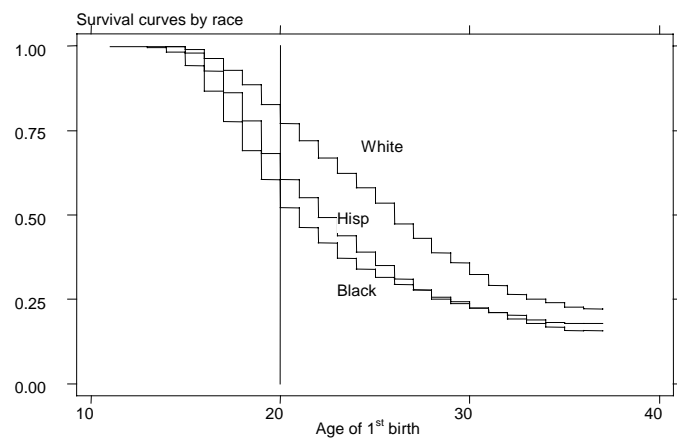


Figure 1. Kaplan-Meier estimates of the proportion of women not having a first birth by a given age. By race and ethnicity (upper left), sample deciles of AFQT (upper right), and late menarche (lower left).

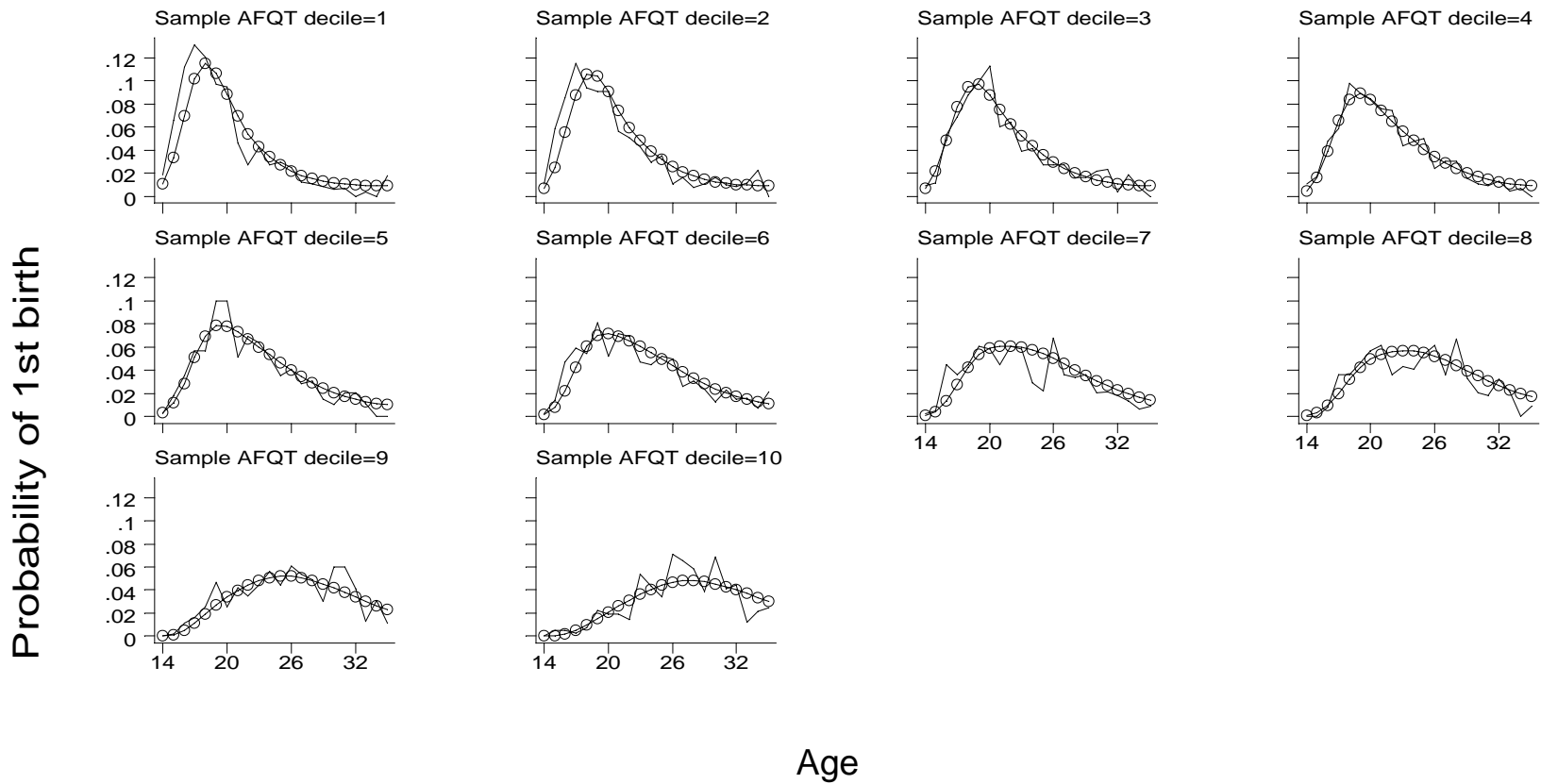


Figure 2a. Comparison of empirical distribution (pdf) of age of 1st birth to estimates from constrained log normal model, with discrete-factor heterogeneity (2 points of support), by sample decile of the AFQT. Constrained log normal estimates derived from model reported in column 4 of Table 6.

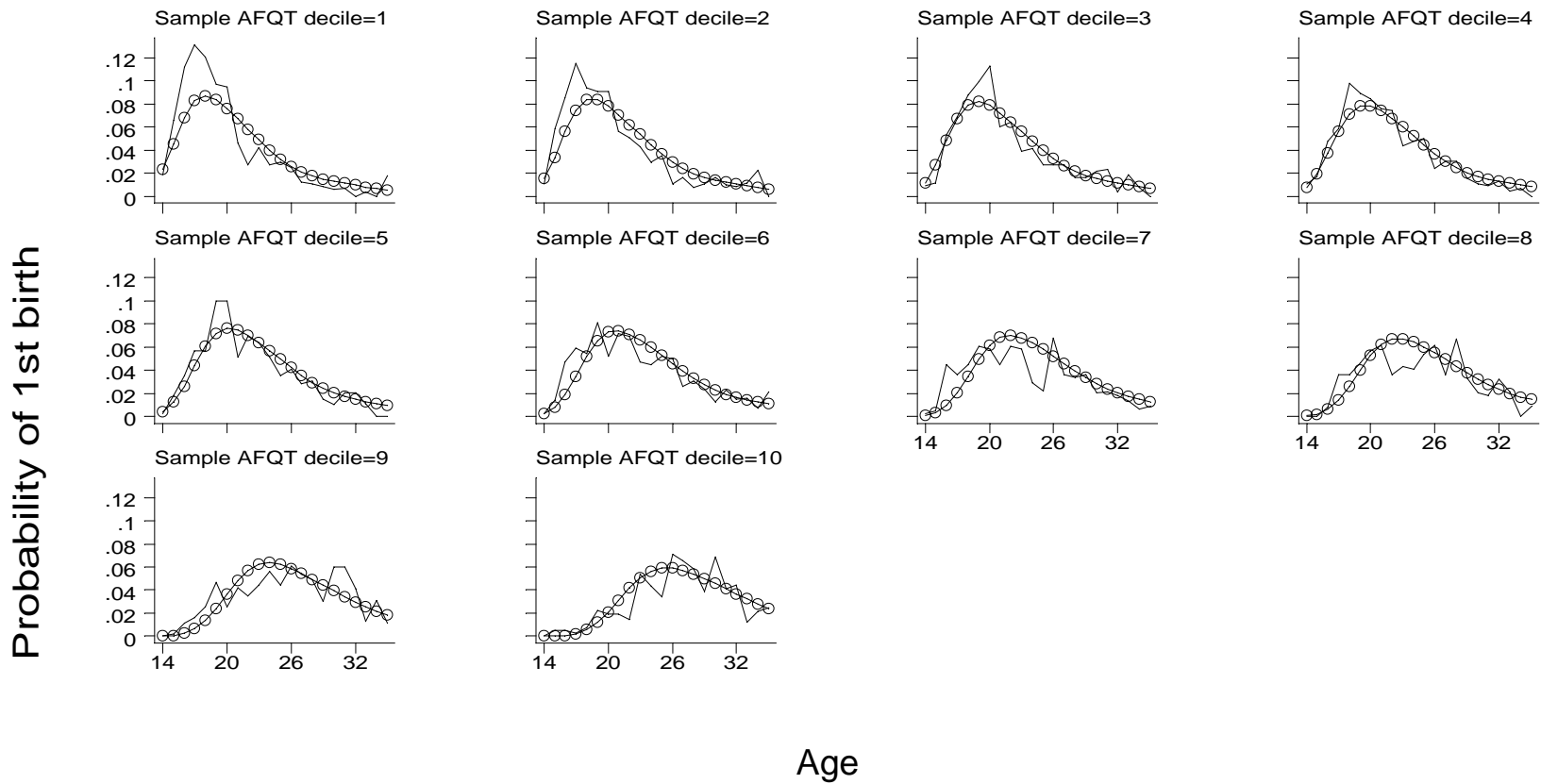


Figure 2b. Comparison of empirical distribution (pdf) of age of 1st birth to estimates from log normal model with discrete heterogeneity (4 points of support), by sample decile of the AFQT. Log normal estimates derived from model reported in column 2 of Table 6.

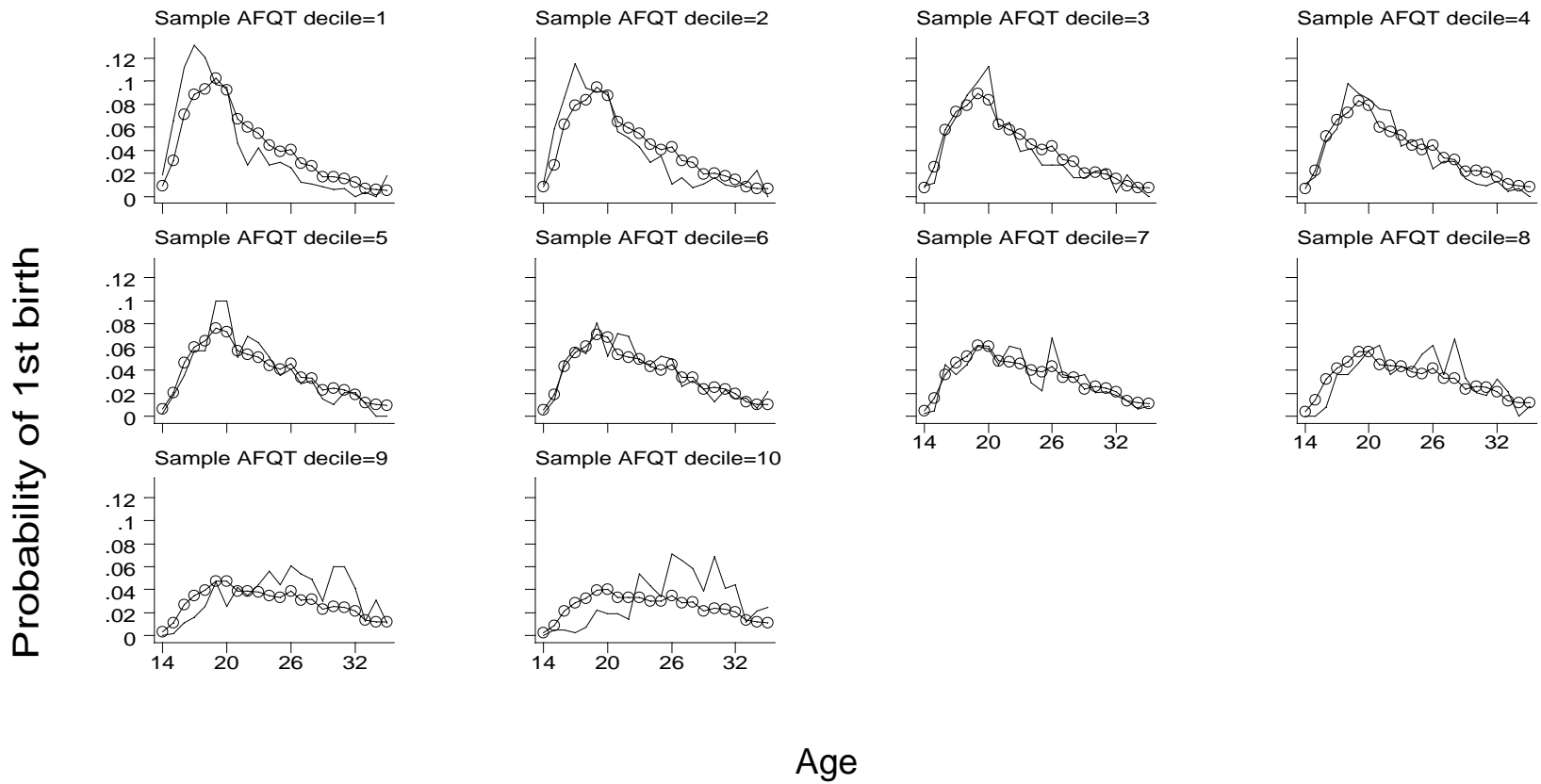


Figure 2c. Comparison of empirical distribution (pdf) of age of 1st birth to estimates from Cox proportional hazard model, by sample decile of the AFQT. Cox estimates derived from model that controls for AFQT percentile, $\ln(\text{age of menarche})$, and the race and family background variables described in Table 1.

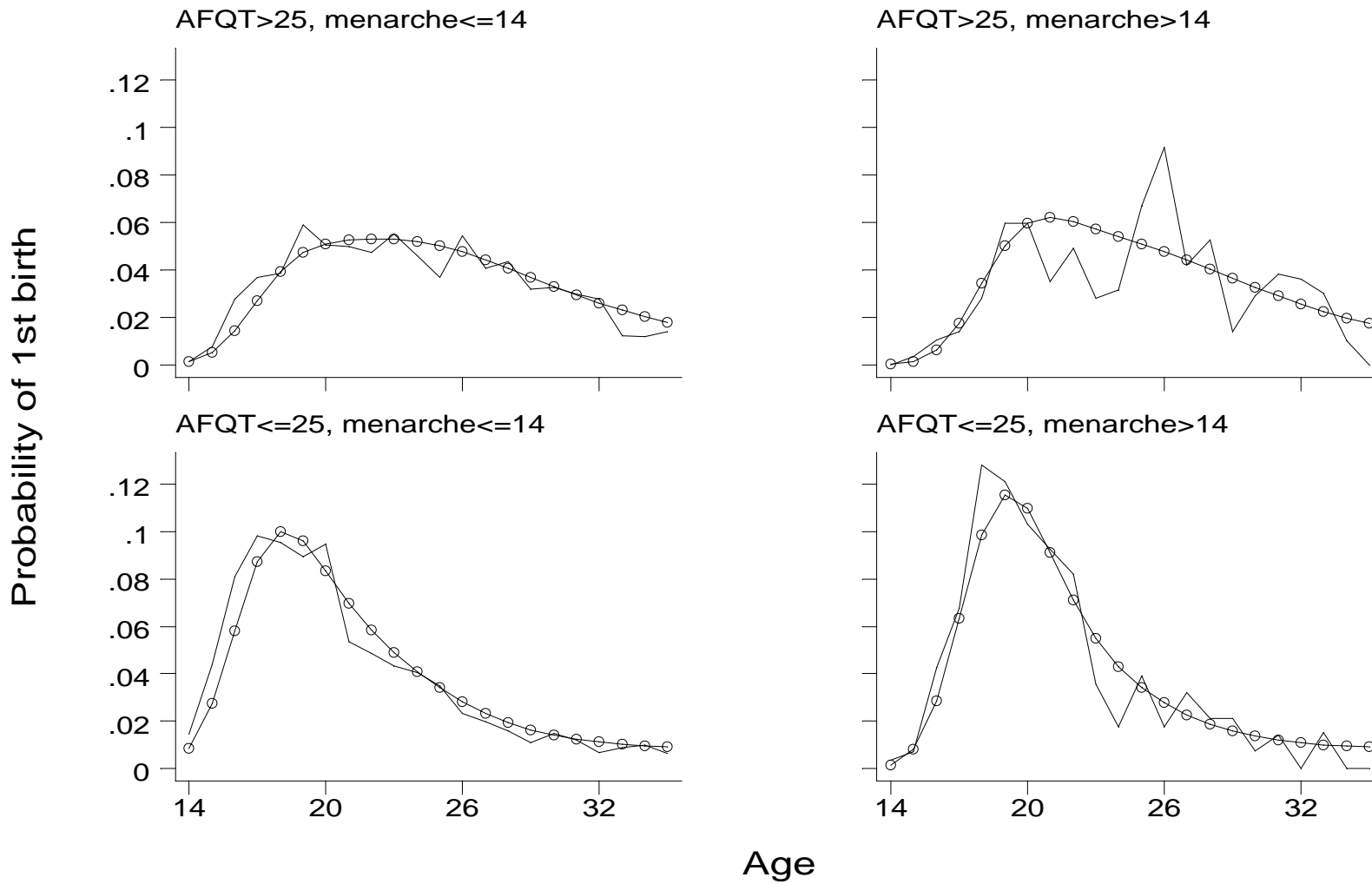


Figure 3a. Comparison of empirical distribution (pdf) of age of 1st birth to estimates from constrained log normal model, with discrete-factor heterogeneity (2 points of support). By AFQT and age of menarche. Constrained log normal estimates derived from model described in Figure 5.

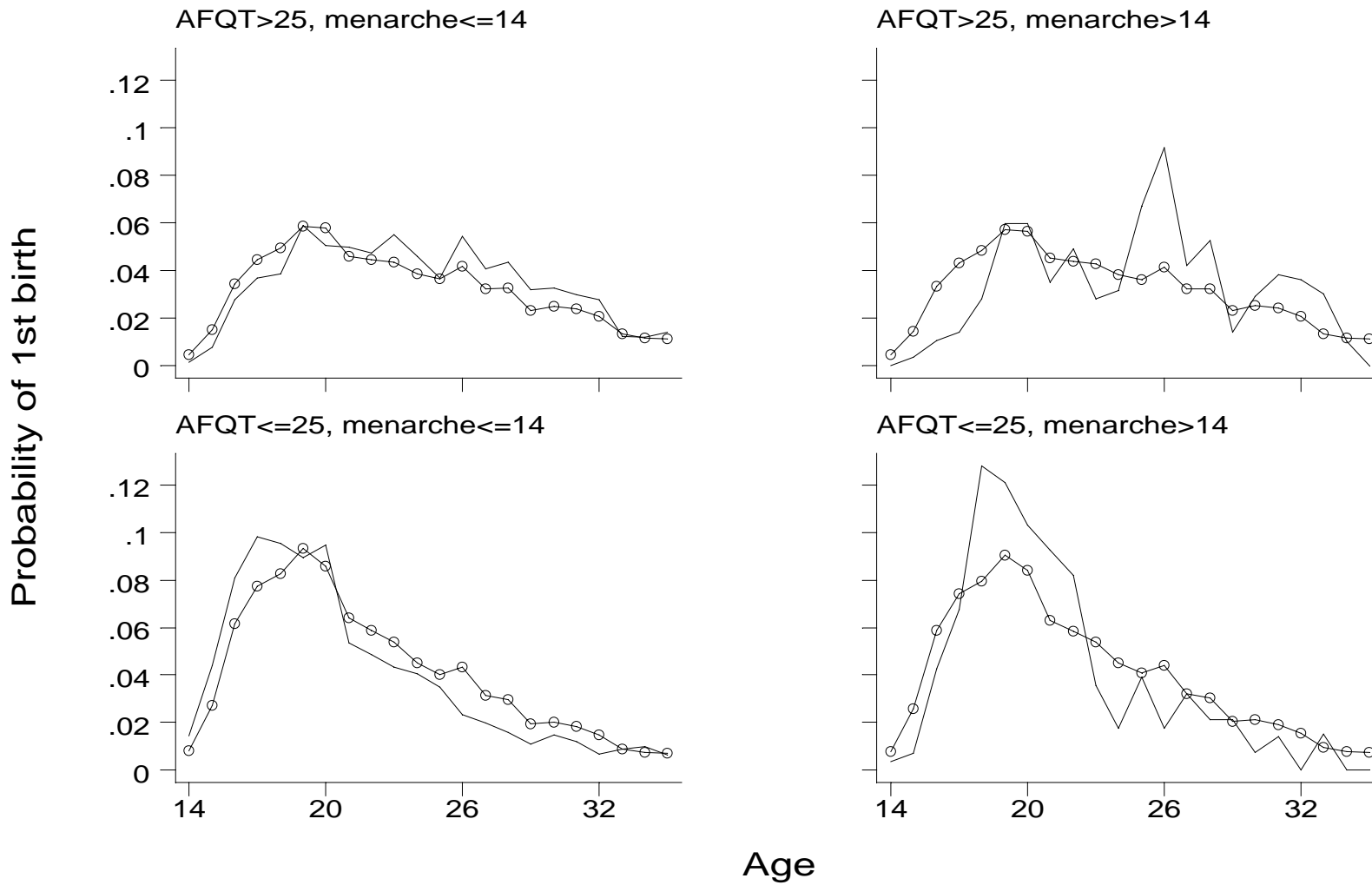


Figure 3c. Comparison of empirical distribution (pdf) of age of 1st birth to estimates from Cox proportional hazard model. By AFQT and age of menarche. Cox estimates derived from model described in Figure 3.

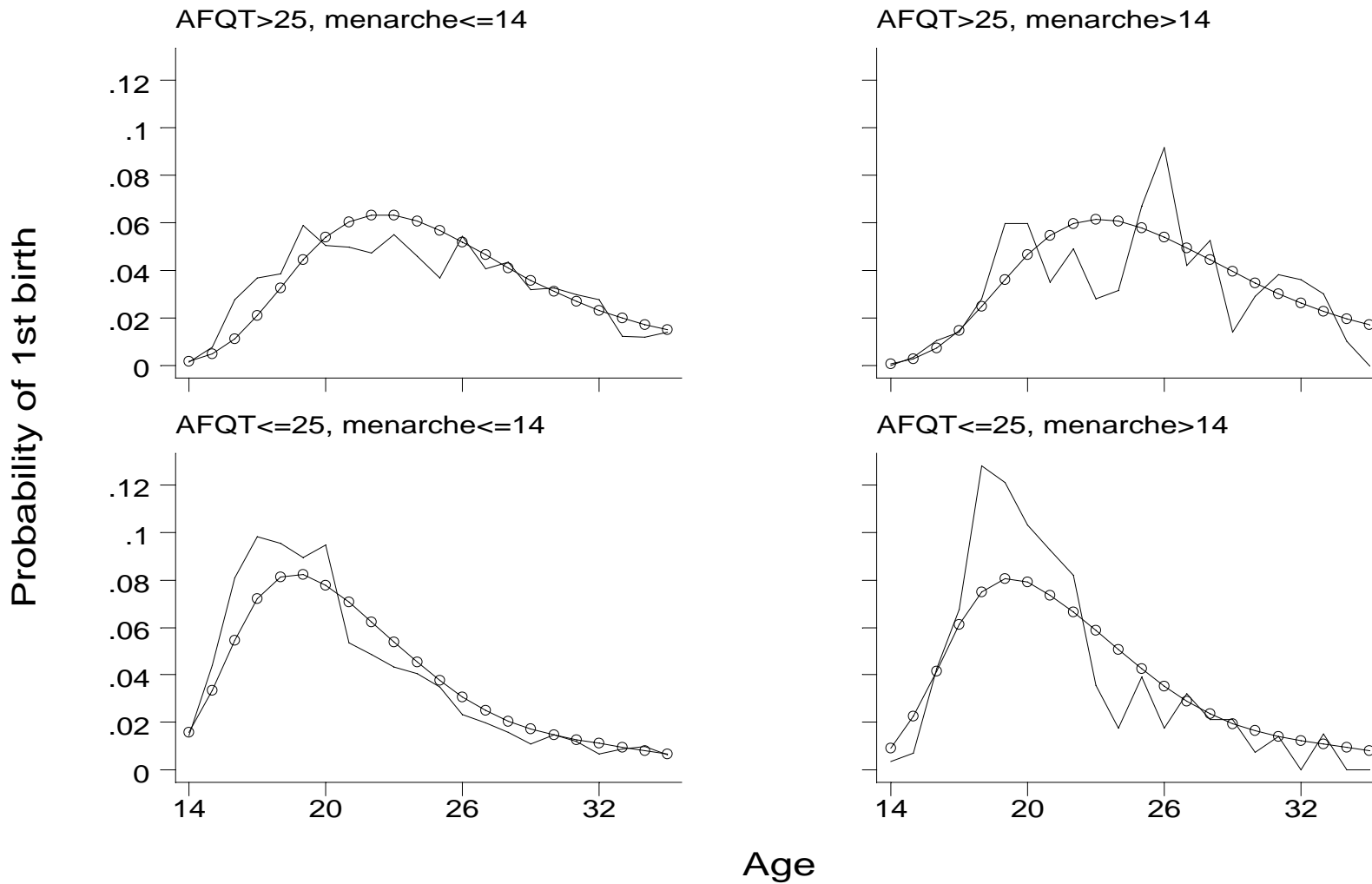


Figure 3b. Comparison of empirical distribution (pdf) of age of 1st birth to estimates from log normal model with discrete heterogeneity (4 points of support). By AFQT and age of menarche. Log normal estimates derived from model described in Figure 4.

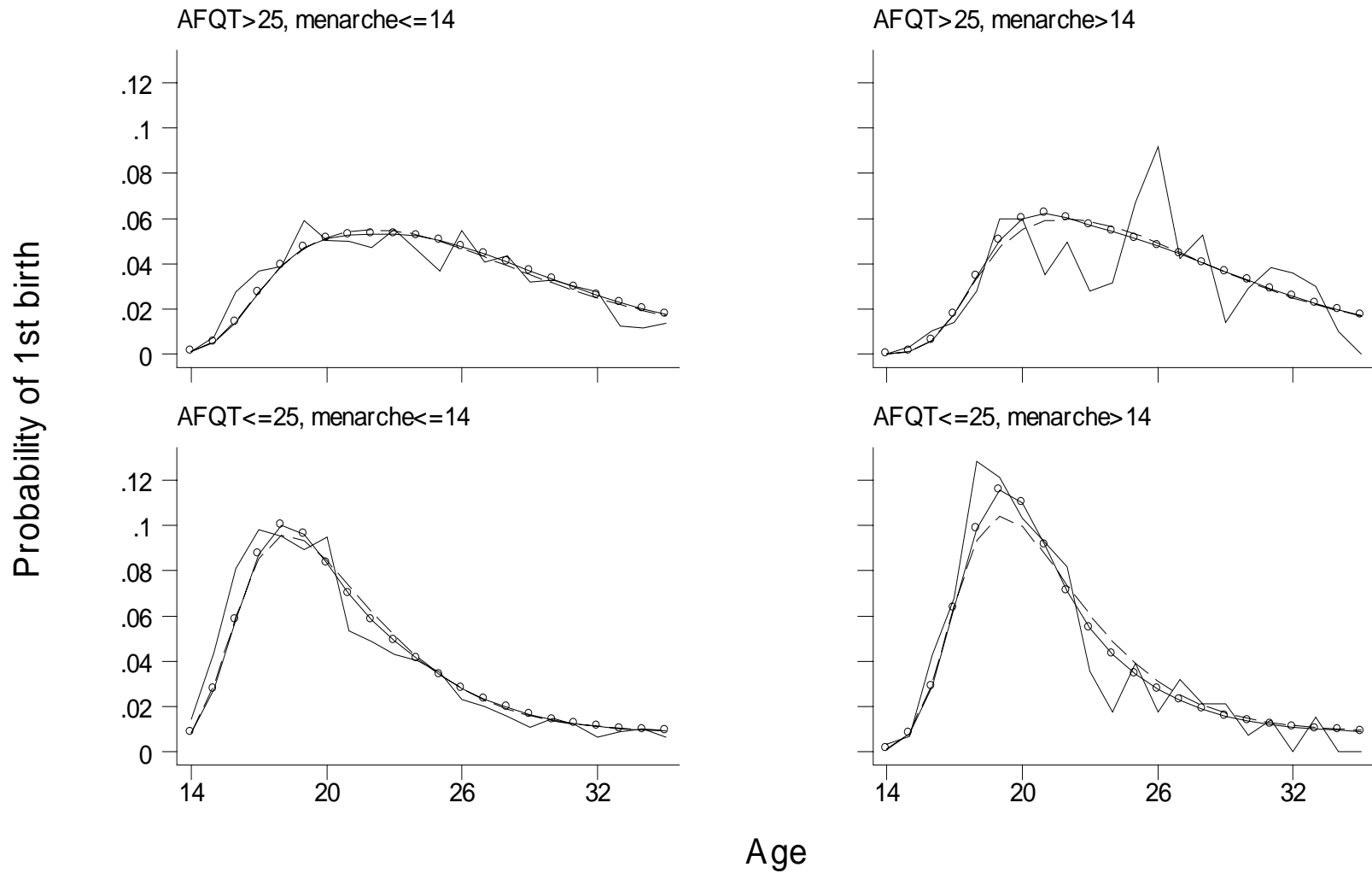


Figure 4. Comparison of empirical distribution (pdf) of age of 1st birth to estimates from constrained log normal (circles) and mistakes (dashed) models with discrete-factor heterogeneity (2 points of support). By AFQT and age of menarche. Constrained log normal and mistake estimates derived from models described in text .

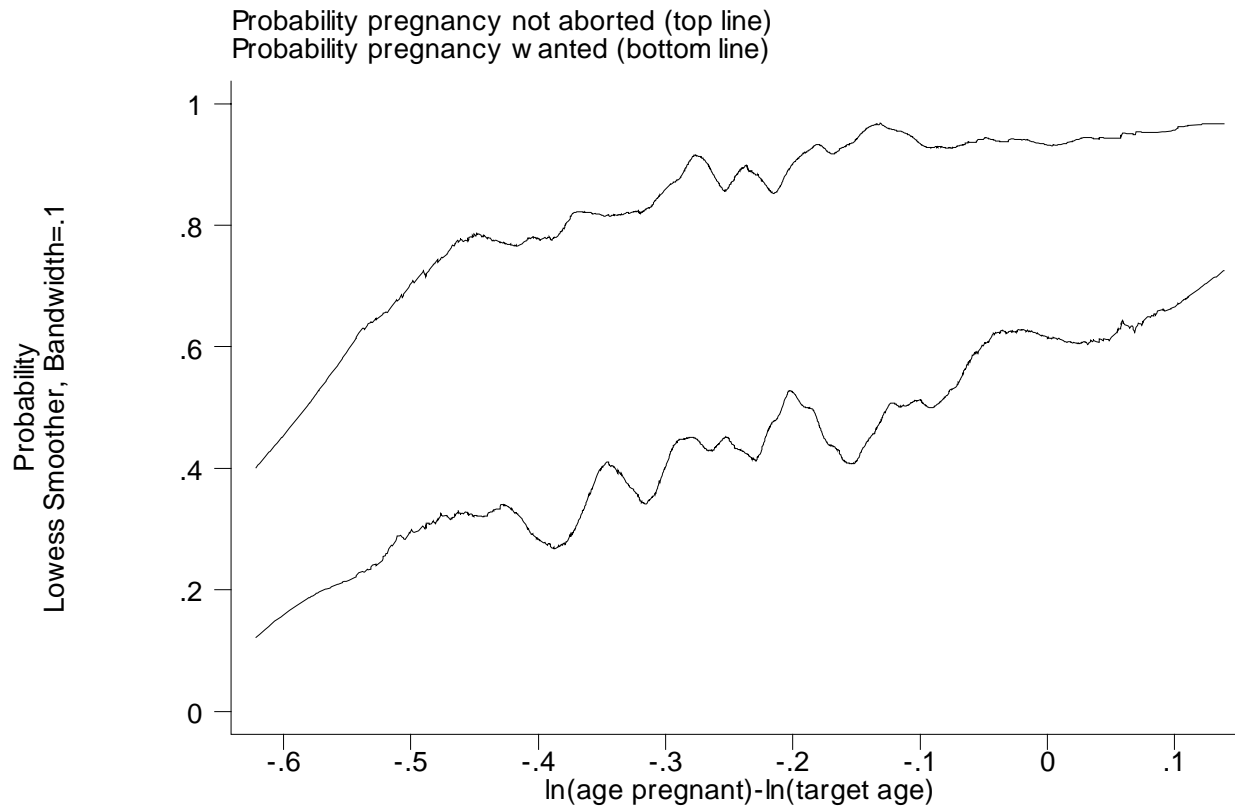


Figure 5. Plot of the probability that first pregnancy was not aborted (top line) or reported to be wanted (bottom line) as a function of the log difference between the age at which the first pregnancy occurred and the woman's target age of first birth (estimated from constrained log normal model). Plotted lines are estimated by non-parametric (Lowess) regression. Sample includes first pregnancies occurring by 1984 (n=1998).

Table 1
Summary statistics by age of menarche

	Total	By Age of first Menstrual Period (Menarche)		
		< 12	12 to 14	> 14
Age of 1st Birth:				
Median	24	23	24	24
% < 18	13.1%	15.8%	13.6%	7.4%
% < 20	26.8%	30.7%	26.4%	24.2%
% < 23	44.2%	47.4%	43.1%	45.3%
% < 29	67.9%	69.4%	67.3%	69.2%
% with no birth by 94	22.9%	23.1%	22.9%	22.7%
Age of Menarche:				
Mean	12.8	10.5	12.8	15.6
Standard Deviation	1.6	0.7	0.7	0.8
% < 12	17.3%	100.0%	0.0%	0.0%
% > 14	13.4%	0.0%	0.0%	100.0%
AFQT Percentile (1-99)				
Mean	38.2	36.7	39.6	33.2
Standard Deviation	27.5	27.1	27.7	26.1
Number of Siblings				
Mean	3.9	3.8	3.8	4.3
Standard Deviation	2.6	2.7	2.6	2.7
Race/Ethnicity				
% Black	30.7%	35.1%	28.3%	37.2%
% Hispanic	19.1%	24.2%	18.6%	15.0%
Mother's Education				
% High School Grad	52.2%	47.1%	53.7%	50.6%
% College Grad	6.8%	6.0%	7.1%	6.5%
Father's Education				
% High School Grad	48.3%	45.1%	49.6%	45.8%
% College Grad	11.4%	10.7%	11.5%	11.3%
Household at Age 14				
% One Parent	32.4%	39.1%	30.8%	45.8%
% with mags/news/lib	11.4%	10.9%	11.4%	12.2%
% with no mags/news/lib	41.5%	39.1%	43.0%	37.2%
Number of Observations	4224	732	2927	565

Table 2

Probit estimates of factors affecting the probability of a teen birth (age 19 or younger)
 Standard errors in parentheses. Probability derivatives in square brackets [dP/dX]

	Full Sample			Age 14 or 15 in 1979		
Black	0.678* (0.048) [0.234]	0.245* (0.055) [0.078]	0.116~ (0.058) [0.036]	0.641* (0.104) [0.219]	0.139 (0.121) [0.043]	0.024 (0.129) [0.007]
Hispanic	0.469* (0.056) [0.164]	0.104 (0.061) [0.033]	-0.064 (0.065) [-0.019]	0.609* (0.117) [0.214]	0.228 (0.129) [0.072]	0.078 (0.135) [0.023]
AFQT Percentile		-0.017* (0.001) [-0.005]	-0.012* (0.001) [-0.004]		-0.021* (0.003) [-0.006]	-0.014* (0.003) [-0.004]
Number of Siblings			0.024* (0.009) [0.007]			0.012 (0.019) [0.003]
One Parent Family at Age 14			0.268* (0.047) [0.083]			0.167 (0.101) [0.049]
Mom Graduated from High School			-0.319* (0.053) [-0.097]			-0.320* (0.115) [-0.092]
Mom Graduated from College			-0.284~ (0.144) [-0.077]			-0.341 (0.352) [-0.086]
Dad Graduated from High School			-0.007 (0.054) [-0.002]			0.018 (0.115) [0.005]
Dad Graduated from College			-0.376* (0.113) [-0.101]			-0.681~ (0.293) [-0.154]
No Magazines, Newspapers or Lib Card at 14			0.051 (0.068) [0.016]			0.128 (0.152) [0.038]
Had Magazines, Newspapers and Lib Card at 14			-0.139* (0.053) [-0.041]			-0.235~ (0.117) [-0.066]
Log Likelihood:	-2349.24	-2192.97	-2113.86	-509.38	-471.43	-452.02
# obs :	4224	4224	4224	924	924	924

Table 3

Impact of late menarche on first birth before age 17 and first birth between ages 18 and 22
 Standard errors in parentheses. Probability derivatives in square brackets [dP/dX]

	Full Sample		AFQT ≤ 25		AFQT > 25	
	By 17	18-22	By 17	18-22	By 17	18-22
Late Menarche	-0.527* (0.091) [-.064]	0.170* (0.060) [.060]	-0.517* (0.108) [-.122]	0.380* (0.083) [.150]	-0.551* (0.172) [-.033]	-0.063 (0.091) [-.018]
Black	0.186* (0.069) [.031]	-0.117~ (0.055) [-.040]	0.098 (0.095) [.027]	0.011 (0.081) [.004]	0.289* (0.103) [.030]	-0.176~ (0.080) [-.050]
Hispanic	-0.167~ (0.081) [-.025]	-0.022 (0.061) [-.007]	-0.237~ (0.109) [-.063]	0.169 (0.091) [.066]	-0.135 (0.126) [-.011]	-0.176~ (0.087) [-.050]
AFQT	-0.012* (0.001) [-.002]	-0.009* (0.001) [-.003]	-0.031* (0.005) [-.009]	0.007 (0.005) [.003]	-0.012* (0.003) [-.001]	-0.013* (0.002) [-.004]
Number of Siblings	0.022~ (0.010) [.003]	0.014 (0.008) [.005]	0.017 (0.013) [.005]	0.012 (0.011) [.005]	0.027 (0.018) [.002]	0.027~ (0.013) [.008]
One Parent Family	0.267* (0.056) [.045]	0.171* (0.045) [.060]	0.258* (0.072) [.073]	0.142~ (0.063) [.055]	0.266* (0.089) [.026]	0.210* (0.066) [.065]
Mom Graduated from High School	-0.291* (0.065) [-.046]	-0.043 (0.051) [-.015]	-0.322* (0.090) [-.085]	0.063 (0.076) [.025]	-0.253* (0.096) [-.024]	-0.139~ (0.069) [-.042]
Mom Graduated from College	-0.336 (0.227) [-.043]	-0.323* (0.112) [-.102]	-0.221 (0.488) [-.056]	0.023 (0.311) [.009]	-0.364 (0.263) [-.025]	-0.334* (0.124) [-.090]
Dad Graduated from High School	-0.021 (0.066) [-.003]	-0.086 (0.051) [-.030]	-0.060 (0.093) [-.017]	-0.067 (0.079) [-.026]	0.006 (0.095) [.0005]	-0.096 (0.068) [-.029]
Dad Graduated from College	-0.456~ (0.180) [-.057]	-0.242* (0.088) [-.079]	-0.036 (0.302) [-.010]	-0.271 (0.231) [-.101]	-0.655* (0.241) [-.040]	-0.172 (0.097) [-.049]
No Magazines, Newspapers, or Lib Card at 14	0.158~ (0.075) [.027]	-0.161~ (0.068) [-.054]	0.165 (0.085) [.048]	-0.084 (0.079) [-.032]	0.07 (0.175) [.006]	-0.104 (0.145) [-.030]
Had Magazines, Newspapers, and Lib Card at 14	-0.132~ (0.067) [-.021]	-0.076 (0.049) [-.026]	-0.055 (0.100) [-.015]	-0.006 (0.083) [-.002]	-0.193~ (0.091) [-.170]	-0.127~ (0.062) [-.038]
Log Likelihood:	-1392.73	-2460.71	-840.29	-1154.86	-542.40	-1279.66
# obs :	4224	4224	461738	1738	2486	2486

Table 4. Maximum likelihood estimates from various models of age of 1st birth
Cox proportional hazard model and lognormal model with and without constraint

	Cox	Log normal		Constrained log normal			
Discrete factor hetero?	no	no	yes	no	yes	no	yes
# points of support:			4		2		2
Coefficients for:	hazard	ln(age of 1st birth)		ln(target age of 1st birth)			
Constant	na	3.16 (0.01)	2.95 (0.01)	3.09 (0.02)	3.04 (0.01)	3.09 (0.02)	3.04 (0.01)
AFQT	-0.0087 (0.0009)	0.0028 (0.0002)	0.0025 (0.0002)	0.0039 (0.0003)	0.0033 (0.0002)	0.0039 (0.0003)	0.0033 (0.0002)
Black	-0.087 (0.047)	0.01 (0.012)	-0.025 (0.009)	0.043 (0.018)	-0.014 (0.015)	0.043 (0.018)	-0.013 (0.014)
Hispanic	-0.117 (0.052)	0.036 (0.013)	0.031 (0.010)	0.052 (0.020)	0.035 (0.016)	0.054 (0.020)	0.037 (0.016)
Ln(age menarche)	-0.223 (0.140)	0.098 (0.036)	0.166 (0.029)	--	--	0.041 (0.055)	0.073 (0.044)
Ln(σ)	na	-1.29 (0.01)	-2.06 (0.04)	-0.98 (0.03)	-1.47 (0.04)	-0.98 (0.03)	-1.47 (0.04)
Coefficients for:				ln(age of biological constraint)			
Constant				2.89 (0.01)	2.81 (0.02)	2.89 (0.01)	2.81 (0.02)
Ln(age menarche)				0.279 (0.041)	0.403 (0.059)	0.269 (0.043)	0.363 (0.065)
Black				-0.081 (0.012)	-0.044 (0.016)	-0.081 (0.012)	-0.044 (0.016)
Hispanic				-0.015 (0.013)	0.02 (0.018)	-0.015 (0.013)	0.019 (0.018)
Ln(σ)				-2.17 (0.04)	-2.29 (0.06)	-2.17 (0.04)	-2.3 (0.06)
Log Likelihood	-25095	-1386	-1043	-1069	-969	-1069	-968

Standard errors are given in parentheses. 4224 observations.

First birth equation in all models also controls for family background variables used in Tables 2 and 3.

Table 5
 Ability of models to predict empirical probability distribution,
 by sample decile of the AFQT

Dependent variable: empirical pdf, by age and AFQT decile
 N=220 (10 AFQT deciles, ages 14-35)

Intercept	-0.001 (0.001)	-0.003 (0.002)	-0.001 (0.002)	-0.003 (0.001)	0.000 (0.001)	-0.003 (0.002)
Predicted Probability Using Constrained Log. Normal Model With Discrete Heterogeneity	1.012* (0.031)			0.799* (0.070)	1.125* (0.126)	0.826 (0.154)
Predicted Probability Using Cox Hazard Model		1.10* (0.045)		0.273* (0.080)		0.269 -0.083
Predicted Probability Using Log Normal Model With Discrete Heterogeneity			1.023* (0.039)		-0.123 (0.132)	-0.026 (0.133)
R-squared	0.826	0.735	0.763	0.835	0.827	0.835

Table 6
 Goodness of fit tests, by AFQT decile and overall
 based on number of births by age (<17,17,18,...,28,29+)

	Constrained Log Normal With Discrete Factor Heterogeneity	Cox Hazard Model	Log Normal Model With Discrete Factor Heterogeneity
By decile			
Decile=1	24.23	42.98	56.66
Decile=2	13.45	29.89	25.25
Decile=3	13.18	25.28	20.54
Decile=4	9.29	24.74	16.32
Decile=5	11.20	21.13	15.29
Decile=6	13.37	13.77	21.40
Decile=7	39.48	16.31	64.15
Decile=8	18.81	39.21	34.69
Decile=9	13.71	44.16	39.39
Decile=10	21.86	97.51	27.88
Overall	178.58	354.96	321.57

Goodness of fit test by decile is chi-squared (d.f.=13) with critical values of 22.36 (5%) and 27.69 (1%).
 Goodness of fit test overall is chi-squared (d.f.=130) with critical values of 157.61 (5%) and 170.42 (1%)

Table 7
Predicting expected age of first birth in 1979

Dependent Variable: ln(expected age of 1st birth)

	Age 14 or 15 in 1979 #obs = 797	Full sample #obs=3049		
Estimated ln(target age) from constrained log normal	0.2417* (0.0276)	0.240* (0.0127)		
AFQT percentile		0.0011* (0.0002)	0.0012* (0.0001)	
Black		0.0298~ (0.0137)	0.0169* (0.0062)	
Hispanic		0.0229 (0.0129)	0.0210* (0.0064)	
Number of siblings		-0.0056~ (0.0022)	-0.0012 (0.0010)	
One parent family at age 14		-0.0256~ (0.0104)	-0.0114~ (0.0051)	
Mom graduated from high school		-0.0063 (0.0110)	0.0084 (0.0054)	
Mom graduated from college		0.0109 (0.0161)	0.0109 (0.0078)	
Dad graduated from high school		0.0016 (0.0111)	0.0040 (0.0054)	
Dad graduated from college		0.0245 (0.0144)	0.0041 (0.0070)	
No magazines, newspapers or library card at age 14		-0.0081 (0.0188)	-0.0142 (0.0084)	
Had magazines, newspapers and library card at age 14		0.0150 (0.0106)	0.0144* (0.0050)	
<hr/>				
R-squared	0.082	0.097	0.101	0.11
p-value from F-test against using target age only		0.271	0.001	

Table 8

Proportion of first pregnancies that were wanted or not aborted, by Age based on sample of all women pregnant by 1984 interview

Age	Number of 1st pregnancies reported	Wanted the pregnancy	Did not abort the pregnancy
10	1	0.00	1.00
12	5	0.40	0.80
13	27	0.26	0.85
14	71	0.20	0.82
15	190	0.33	0.87
16	269	0.35	0.88
17	282	0.33	0.85
18	320	0.47	0.87
19	283	0.49	0.89
20	165	0.58	0.87
21	139	0.63	0.91
22	106	0.58	0.89
23	71	0.65	0.83
24	30	0.77	0.90
25	23	0.70	0.91
26	7	0.57	1.00
Total	1,989	0.45	0.87

Table 9
Predicting whether first pregnancy was wanted or not aborted

Dependent variable:	Pregnancy was wanted		Pregnancy was not aborted	
Ln(age preg + 1) minus estimated ln(target age)	0.6703*	0.8213*	0.5199*	0.4845*
	(0.0629)	(0.0651)	(0.0510)	(0.0529)
Black		-0.2307*	-0.1823*	0.0426~
		(0.0247)	(0.0278)	(0.0167)
Hispanic		0.0339	0.0333	0.0244
		(0.0294)	(0.0317)	(0.0203)
Ln(age preg + 1)			0.9860*	0.3150*
			(0.0814)	(0.0590)
AFQT percentile			-0.0026*	-0.0022*
			(0.0006)	(0.0004)
Number of siblings			0.0014	0.0064~
			(0.0041)	(0.0025)
One parent family at age 14			0.0072	0.0070
			(0.0227)	(0.0151)
Mom graduated from high school			-0.0244	-0.0269
			(0.0255)	(0.0180)
Mom graduated from college			0.0408	-0.1110
			(0.0634)	(0.0609)
Dad graduated from high school			-0.0204	-0.0122
			(0.0260)	(0.0184)
Dad graduated from college			-0.0872	-0.1422*
			(0.0512)	(0.0494)
No magazines, newspapers or library card at age 14			0.0598	0.0098
			(0.0318)	(0.0171)
Had magazines, newspapers and library card at age 14			-0.0292	-0.0274
			(0.0256)	(0.0191)
R-squared	0.048	0.101	0.111	0.063
				0.066
				0.089
P-value of F-test against model in previous column		0.000	0.009	0.038
				0.000

OLS regressions. Robust Standard errors given in parentheses.