The effects of teenage fertility on young adult childbearing[*]

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Abstract:
Numerous studies of fertility behavior find that an early age at first birth increases the rate of subsequent childbearing. Typically, however, these studies do not account for the possibility of serial correlation in the unobserved determinants of fertility. Using 1979 — 1992 individual-level data from the National Longitudinal Survey of Youth, this paper employs the Method of Simulated Moments to estimate panel probit models of annual birth outcomes. The panel probit models account for several alternative sources of serial correlation. Estimation reveals that once serial correlation is taken into account, the subsequent fertility effects of early childbearing are either statistically eliminated or reversed.
JEL classification: J13
Key words: Fertility, method of simulated moments

Article:
1. Introduction
Social scientists have devoted substantial effort to studying the causes and consequences of teenage childbearing in the United States. Adolescent fertility has also become a focus of policy with several recent proposals including President Clinton's and congressional Republicans' welfare reform plans aimed at reducing out-of-wedlock births among teenagers. Academic and public concern rests in large part with the negative economic and social outcomes associated with early childbearing and the possibility that some outcomes, such as teen mothers' high rates of subsequent fertility, may exacerbate others.

Studies that have examined the effects of early childbearing on subsequent fertility have generally concluded that giving birth at an early age increases the rate of subsequent childbearing. In her review of this literature, Hofferth (1987) writes

> All the evidence supports the conclusion that early child bearers have more children, especially more unwanted children, and that they have them more rapidly than older child bearers.

These findings have been replicated numerous times using data from several countries. Nevertheless, there remains some doubt as to whether early fertility actually affects later childbearing.

This doubt arises because studies have typically failed to account for the possibility of serial correlation in the unobserved determinants of fertility. It is well known that ordinary regression estimates of the effects of past on current behavior may be biased in the presence of serially correlated errors. In the specific case of fertility, Heckman et al. (1985) and Heckman and Walker (1991) have demonstrated, using data on Swedish women, that serial correlation can lead to a spurious positive estimated relationship between early births and subsequent

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childbearing. More generally, several other recent studies have documented how sensitive estimates of various consequences of early fertility are to alternative statistical controls for biases from unobserved heterogeneity.[1]

This paper examines the relationship between teenage and young adult childbearing using 1979 — 1992 data from the National Longitudinal Survey of Youth (NLSY). Specifically, it applies the Method of Simulated Moments (MSM) panel probit procedure developed by Keane (1992, 1994) to estimate the determinants of annual fertility from ages 15 to 28 separately for white, black, and Hispanic women. The MSM procedure is advantageous because it provides a tractable way of consistently estimating dynamic probit models with general forms of serial correlation.

The paper develops a number of important results. First, it finds that the unobserved determinants in several specifications of the fertility model are characterized by moderately complicated forms of serial correlation. Thus, the use of the MSM estimator, rather than standard cross-section or random effects probit procedures, is justified. Second and more substantively, it finds that controlling for serial correlation makes a difference in the estimated relationship between early and subsequent fertility. Models that do not account for serial correlation generally reproduce the literature's findings of a strong positive association between early and later births for white, black, and Hispanic women. However, after serial correlation is taken into account, this association reverses direction for white and black women; for Hispanic women, the association diminishes in magnitude and becomes statistically insignificant.

The remainder of this paper is organized as follows. The next section briefly discusses the key hypotheses and methodological issues and reviews the findings of previous studies. Section 3 describes the data that are used. The paper's empirical results are reported and interpreted in Sect. 4. Concluding remarks appear in Sect. 5.

2. Research issues

Dynamic economic models of life-cycle fertility are consistent with the possibility of both actual effects of early births on subsequent childbearing and "spurious" effects through omitted and unobserved variables.[2] These models are generally extensions of household production models in which period-by-period fertility enters as both a good and a decision variable and previous fertility (or more precisely, the time and goods requirements of children) enters as an argument in the production of household commodities. In such a framework, previous births may have direct effects on subsequent childbearing if, for instance, women have preferences regarding the number of children and use fertility as a stock adjustment variable. Previous fertility may also indirectly influence later childbearing through its effects on human capital investment, productivity inside and outside the labor market, and marriage behavior. For example, a negative effect of teen fertility on human capital investment might affect subsequent market and non- market opportunities thereby altering the incentives and resources available for later childbearing.[3] Negative effects of early childbearing on subsequent marriage prospects and marital stability might similarly reduce the resources available for later childbearing. Finally, while direct and indirect effects from early to subsequent childbearing are possible, the dynamic models clearly demonstrate that causality could also run from preferences for children and anticipated family size to fertility timing.

The empirical relationship between birth timing and subsequent fertility has been extensively researched. As mentioned, the early literature reported a positive relationship between early births and subsequent childbearing but mostly failed to control for sources of spurious correlation. The most prominent studies addressing the problem of unobserved heterogeneity were the analyses of Swedish women by Heckman et al. (1985) and Heckman and Walker (1991). These studies showed that when serial correlation was taken into account, the positive association between early and later births was either eliminated or reversed. Given these results and the high rates of adolescent childbearing in the United States, it is surprising that longitudinal models incorporating detailed controls for serial correlation have not been used in a general analysis of the relationship between teenage and subsequent fertility in this country.

Several studies have examined selected groups of U.S. women or included limited or specialized correlation controls. Hotz and Miller (1988) jointly modeled married women's longitudinal contraception decisions, labor supply, wages, and child care costs and included controls for unobserved heterogeneity. Their results indicated that an earlier age at birth decreased subsequent birth probabilities. In contrast, Hoffman et al. (1993) found that the positive effects of teen fertility on subsequent childbearing were reduced but not eliminated when sibling comparisons were used to control for unobserved heterogeneity. Using miscarriages as a quasi-experimental control for heterogeneity, Hotz et al. (1995) came to a similar conclusion.[4]

This paper examines the determinants of young women's fertility using a simple longitudinal econometric specification. Let $y_{i,t}$ be an observed dichotomous variable that equals one if woman i experiences a birth at age t and zero otherwise, and let $y^*_{i,t}$ be an unobserved latent variable that describes the woman's propensity to give birth. The latent variable is assumed to be a function of previous fertility behavior, other observed and possibly time-varying determinants ($x_{i,t}$), and unobserved factors ($\varepsilon_{i,t}$) such that

$$y^*_{i,t} = A(L)y_{i,t}\beta_y + x_{i,t}\beta_x + \varepsilon_{i,t} \quad \text{where} \quad i = 1, N \ , \quad t = 1, T \ , \tag{1}$$

and A (L) is a general lag operator. The observed birth outcomes at each age are assumed to follow

$$y_{i,t} = \begin{cases} 1 & \text{if} \quad y^*_{i,t} > 0 \\ 0 & \text{otherwise} \end{cases} . \tag{2}$$

With the additional assumption that e4, is independently and identically normally distributed, Eqs. (1, 2) constitute a simple probit specification.

If the explanatory variables are uncorrelated with the unobserved component, the coefficients in Eq. (1) can be consistently estimated using standard maximum likelihood (ML) methods. However, in the presence of serially correlated errors, this orthogonality condition breaks down for the lagged dependent variables, and consistent estimation becomes problematic. While ML estimation is theoretically possible, it requires the evaluation of T-fold integrals for general forms of serial correlation.[5] In the case of normally distributed variables and data sets with T> 3, standard numerical evaluation methods are impractical. Fortunately, newly-refined statistical techniques based on simulation methods provide alternatives to standard ML estimation.

This paper adopts the MSM panel probit procedure developed by Keane (1992, 1994); a detailed description of this procedure appears in Appendix A. Keane's estimator is advantageous in several respects. First, it permits consistent estimation of dynamic probit panel models from long panels with complex correlation structures performing well even in the presence of high degrees of serial correlation. Second, given a consistent set of parameter starting values, the estimator is asymptotically efficient. Third, consistency and asymptotic normality in the sample size hold for fixed numbers of simulations; thus, estimation can be carried out in a reasonable amount of time on a personal computer or small workstation.

3. Data

The primary data used in this analysis come from the 1979— 1992 panels of the National Longitudinal Survey of Youth (Center for Human Resource Research 1992). The NLSY is a large national sample of 12686 individuals who were 14 to 21 years old in 1979.[6] Individuals have been re-interviewed annually since 1979. The survey contains detailed longitudinal demographic and geographic information. Personal and family background data are also available. Attrition in the sample has been minimal.

The empirical analysis requires year-to-year data on teenage and young adult childbearing as well as contemporaneous data on the determinants of childbearing. Accordingly, the analysis focuses on women who were 14-16 years old on January 1, 1979. For each woman, the paper constructs an annual fertility history describing whether she experienced a birth in each calender year from the time she was age 15 until she either left the survey, was missing information regarding her childbearing or one of the explanatory variables, or reached the 1992 panel. The annual binary fertility outcomes serve as the dependent variable in the empirical analysis. The data form an unbalanced panel that contains up to 14 annual observations for each woman. Women who experienced births prior to the calendar year in which they turned 16 (the calendar year they entered as 15-year-olds) are excluded from the analysis.[7] Separate data sets are generated for black, Hispanic, and other (mostly white and hereafter referred to as white/other) women.

To examine the effect that a woman's prior fertility has on her subsequent childbearing, the analysis uses distributed lags in the dichotomous birth measure as explanatory variables. Specifically, the models include one-, two-, three-, and four-year lags in the fertility variable as well as a variable which is the sum of all births lagged five years or more. [8] To capture differences in the effects of first and subsequent births, several models include similar lag variables defined only over higher parity births. Finally, to examine the effects of early childbearing, the models include lags defined only over births and occurred prior to the calendar year in which the woman turned 20 years of age.

Beyond the lags in the dependent variable, the empirical specifications include a modest number of independent controls.[9] Several family background measures — the woman's number of siblings, an indicator for whether she lived in a non- intact family at age 14, an indicator for the availability of reading materials at age 14, and three variables for her mother's education — act as time invariant controls. Time-varying measures directly available from the NLSY include the woman's age, an indicator for residence in an urban area, and the unemployment rate in her country of residence. In addition, annual county-specific information on birth rates (U.S. Bureau of Health Professions 1993), the number of obstetricians and gynecologists (Ob-Gyn physicians) (data obtained from Quality Resource Systems, Inc. of Fairfax, Virginia), and average earnings for retail workers (U.S. Bureau of Economic Analysis 1992) as well as annual state-level information on Medicaid abortion funding restrictions (Merz 1994; Matthews et al. 1995) have been merged into the data using geographic codes from the NLSY. All of the independent explanatory variables have appeared in previous analyses of early fertility. Each has a straightforward economic, sociological, or demographic interpretation.

The final data set contains 11933 annual observations describing 1128 individual white/other women, 5931 annual observations describing 531 individual black women, and 4118 annual observations describing 384 individual Hispanic women. Means and standard deviations for the dependent and independent variables calculated separately by race and ethnicity appear in Appendix B. The number of observations in each race/age cohort of the panel are also listed in Appendix B.

4. Empirical analysis

Estimation results from alternative specifications of the panel probit models for white/other, black, and Hispanic women appear in Tables 1, 2, and 3, respectively. The first three columns in each table report results from an initial set of specifications which pool the longitudinal fertility data for each woman as independent observations and estimate the determinants of year-to-year childbearing using

Table 1. Probit analysis of the effects of teenage fertility on subsequent childbearing using alternative serial correlation and birth history controls – Non-black, Non-hispanic (white/other) women

	Alternative controls for birth history/ no controls for serial correlation			No controls for birth history/ alt. controls for serial correlation			Alternative controls for birth history and serial correlation		
	(a)	(b)	(c)	(d)	(e)	(f)	(g)	(h)	(i)
Lagged birth variables									
Any births in previous year?	–	−0.244[c] (0.075)	0.045 (0.095)	–	–	–	−0.596[c] (0.047)	−0.559[c] (0.058)	−0.900[c] (0.178)
Any births two years ago?	–	0.289[c] (0.068)	0.655[c] (0.087)	–	–	–	−0.325[c] (0.050)	−0.264[b] (0.103)	−0.673[c] (0.224)
Any births three years ago?	–	0.233[c] (0.077)	0.552[c] (0.099)	–	–	–	−0.359[c] (0.053)	−0.342[c] (0.114)	−0.576[a] (0.341)
Any births four years ago?	–	0.002 (0.091)	0.424[c] (0.115)	–	–	–	−0.510[c] (0.061)	−0.471[c] (0.121)	−0.601[b] (0.306)
All births five or more years ago	–	−0.249[c] (0.071)	0.085 (0.098)	–	–	–	−0.718[c] (0.056)	−0.728[c] (0.124)	−0.687[c] (0.177)
Lagged higher parity birth variables									
Higher parity birth in previous year?	–	–	0.704[c] (0.139)	–	–	–	–	−0.071 (0.128)	–
Higher parity birth two years ago?	–	–	−0.893[c] (0.133)	–	–	–	–	−0.164 (0.161)	–
Higher parity birth three years ago?	–	–	−0.631[c] (0.149)	–	–	–	–	−0.022 (0.170)	–
Higher parity birth four years ago?	–	–	−0.713[c] (0.183)	–	–	–	–	−0.031 (0.195)	–
Higher parity births five or more years ago	–	–	−0.395[c] (0.147)	–	–	–	–	0.099 (0.190)	–

Lagged early birth variables									
Teen birth in previous year?	0.226[b]	0.472[c]	0.292[b]	0.099	0.342[c]	0.310[b]	−0.116	−0.157[a]	−0.023
	(0.114)	(0.135)	(0.141)	(0.115)	(0.132)	(0.129)	(0.096)	(0.094)	(0.101)
Teen birth two years ago?	0.315[c]	0.049	−0.129	0.192	0.031	−0.091	−0.326[c]	−0.377[c]	−0.157[a]
	(0.109)	(0.127)	(0.132)	(0.119)	(0.127)	(0.130)	(0.088)	(0.086)	(0.088)
Teen birth three years ago?	0.432[c]	0.206	0.131	0.314[c]	0.221[b]	0.393[c]	−0.210[b]	−0.229[b]	−0.105
	(0.106)	(0.129)	(0.135)	(0.111)	(0.112)	(0.119)	(0.090)	(0.090)	(0.089)
Teen birth four years ago?	0.248[b]	0.263[a]	0.173	0.133	0.049	0.118	−0.138	−0.168	−0.055
	(0.112)	(0.143)	(0.150)	(0.122)	(0.122)	(0.125)	(0.109)	(0.108)	(0.108)
Teen births five or more years ago	−0.212[c]	0.034	0.021	−0.331[c]	−0.404[c]	−0.309[c]	−0.328[c]	−0.349[c]	−0.304
	(0.056)	(0.093)	(0.093)	(0.072)	(0.074)	(0.084)	(0.077)	(0.081)	(0.190)
Covariance parameters									
σ_{μ}^{2}	–	–	–	0.062[c]	0.105[c]	0.052[a]	0.701[c]	0.733[c]	0.591[a]
				(0.021)	(0.023)	(0.031)	(0.056)	(0.096)	(0.358)
$\rho(1)$	–	–	–	–	−0.189[c]	−0.087[a]	–	–	0.763
					(0.046)	(0.048)			(0.567)
$\rho(2)$	–	–	–	–	–	0.152[c]	–	–	–
						(0.054)			
Simulated log likelihood	−3187.82	−3163.48	−3126.24	−3184.45	−3172.94	−3171.83	−3113.25	−3111.26	−3119.62

Note: Estimates based on 1128 individual (11933 person-year) observations from the 1979 – 1992 panels of the NLSY. Coefficients for number of siblings, non-intact family at age 14, availability of reading materials at age 14, mother's education (three variables), age, age squared, local birth rate, availability of Ob-Gyn physicians, restrictions on Medicaid funding of abortions, urban residence, local unemployment rate, average annual retail earnings, and an intercept are estimated but not reported. Standard errors appear in parentheses.

[a] Significant at 0.10 level.

[b] Significant at 0.05 level.

[c] Significant at 0.01 level.

Table 2. Probit analysis of the effects of teenage fertility on subsequent childbearing using alternative serial correlation and birth history controls – Black women

	Alternative controls for birth history/ no controls for serial correlation			No controls for birth history/ alt. controls for serial correlation			Alternative controls for birth history and serial correlation	
	(a)	(b)	(c)	(d)	(e)	(f)	(g)	(h)
Lagged birth variables								
Any births in previous year?	–	−0.191[b] (0.087)	−0.133 (0.133)	–	–	–	−0.493[c] (0.072)	−0.447[b] (0.195)
Any births two years ago?	–	0.415[c] (0.079)	0.532[c] (0.113)	–	–	–	−0.035 (0.097)	0.147 (0.259)
Any births three years ago?	–	0.188[b] (0.090)	0.374[c] (0.128)	–	–	–	−0.194[b] (0.086)	0.008 (0.258)
Any births four years ago?	–	0.029 (0.102)	0.166 (0.139)	–	–	–	−0.308[c] (0.094)	−0.173 (0.263)
All births five or more years ago	–	−0.014 (0.072)	0.224[b] (0.110)	–	–	–	−0.346[c] (0.074)	−0.133 (0.255)
Lagged higher parity birth variables								
Higher parity birth in previous year?	–	–	−0.100 (0.156)	–	–	–	–	−0.008 (0.184)
Higher parity birth two years ago?	–	–	−0.206 (0.141)	–	–	–	–	−0.115 (0.181)
Higher parity birth three years ago?	–	–	−0.310[a] (0.161)	–	–	–	–	−0.195 (0.219)
Higher parity birth four years ago?	–	–	−0.180 (0.174)	–	–	–	–	−0.077 (0.222)
Higher parity births five or more years ago	–	–	−0.320[c] (0.122)	–	–	–	–	−0.193 (0.195)

Lagged early birth variables								
Teen birth in previous year?	−0.027	0.168	0.137	−0.207	0.027	0.010	−0.179	−0.055
	(0.123)	(0.148)	(0.161)	(0.130)	(0.149)	(0.143)	(0.144)	(0.183)
Teen birth two years ago?	0.432[c]	0.037	−0.023	0.257[b]	0.064	0.004	−0.230[a]	−0.149
	(0.107)	(0.130)	(0.140)	(0.115)	(0.126)	(0.130)	(0.125)	(0.170)
Teen birth three years ago?	−0.042	−0.180	−0.265[a]	−0.200	−0.289[b]	−0.185	−0.382[c]	−0.343[b]
	(0.121)	(0.148)	(0.158)	(0.124)	(0.123)	(0.135)	(0.124)	(0.156)
Teen birth four years ago?	0.011	−0.045	−0.077	−0.148	−0.226[a]	−0.190	−0.288[b]	−0.185
	(0.122)	(0.156)	(0.164)	(0.124)	(0.125)	(0.126)	(0.136)	(0.181)
Teen births five or more years ago	0.027	0.029	0.0004	−0.143[b]	−0.225[c]	−0.167[b]	−0.216[a]	−0.114
	(0.049)	(0.088)	(0.089)	(0.067)	(0.070)	(0.075)	(0.112)	(0.153)
Covariance parameters								
σ_μ^2	–	–	–	0.099[c]	0.154[c]	0.117[c]	0.464[c]	0.283
				(0.032)	(0.036)	(0.038)	(0.106)	(0.209)
$\varrho(1)$	–	–	–	–	−0.222[c]	−0.145[b]	–	–
					(0.050)	(0.058)		
$\varrho(2)$	–	–	–	–	–	0.108[a]	–	–
						(0.060)		
Simulated log likelihood	−2043.55	−2025.69	−2021.00	−2038.71	−2028.82	−2026.61	−2020.03	−2017.82

Note: Estimates based on 531 individual (5931 person-year) observations from the 1979 – 1992 panels of the NLSY. Coefficients for number of siblings, non-intact family at age 14, availability of reading materials at age 14, mother's education (three variables), age, age squared, local birth rate, availability of Ob-Gyn physicians, restrictions on Medicaid funding of abortions, urban residence, local unemployment rate, average annual retail earnings, and an intercept are estimated but not reported. Standard errors appear in parentheses.

[a] Significant at 0.10 level.

[b] Significant at 0.05 level.

[c] Significant at 0.01 level.

Table 3. Probit analysis of the effects of teenage fertility on subsequent childbearing using alternative serial correlation and birth history controls – Hispanic women

	Alternative controls for birth history/ no controls for serial correlation			No controls for birth history/ alt. controls for serial correlation			Alternative controls for birth history and serial correlation	
	(a)	(b)	(c)	(d)	(e)	(f)	(g)	(h)
Lagged birth variables								
Any births in previous year?	–	−0.159 (0.110)	−0.092 (0.162)	–	–	–	−0.313 (0.243)	−0.457[b] (0.227)
Any births two years ago?	–	0.183[a] (0.106)	0.384[c] (0.144)	–	–	–	0.012 (0.274)	−0.033 (0.320)
Any births three years ago?	–	0.112 (0.121)	0.290[a] (0.162)	–	–	–	−0.057 (0.295)	−0.116 (0.315
Any births four years ago?	–	0.303[b] (0.128)	0.388[b] (0.173)	–	–	–	0.132 (0.313)	−0.039 (0.366)
All births five or more years ago	–	−0.155 (0.110)	−0.038 (0.156)	–	–	–	−0.305 (0.251)	−0.432 (0.327)
Lagged higher parity birth variables								
Higher parity birth in previous year?	–	–	−0.124 (0.197)	–	–	–	–	0.029 (0.215)
Higher parity birth two years ago?	–	–	−0.386[b] (0.191)	–	–	–	–	−0.221 (0.260)
Higher parity birth three years ago?	–	–	−0.317 (0.216)	–	–	–	–	−0.176 (0.276)
Higher parity birth four years ago?	–	–	−0.090 (0.227)	–	–	–	–	0.030 (0.270)
Higher parity births five or more years ago	–	–	−0.117 (0.189)	–	–	–	–	0.028 (0.274)

Lagged early birth variables								
Teen birth in previous year?	0.257[a]	0.417[b]	0.369[a]	0.176	0.378[b]	0.380[b]	0.362	0.212
	(0.147)	(0.181)	(0.199)	(0.150)	(0.181)	(0.178)	(0.221)	(0.250)
Teen birth two years ago?	0.505[c]	0.338[b]	0.212	0.427[c]	0.284[a]	0.276	0.316	0.100
	(0.138)	(0.172)	(0.184)	(0.162)	(0.166)	(0.182)	(0.204)	(0.211)
Teen birth three years ago?	0.138	0.047	−0.014	0.067	−0.013	0.018	0.060	−0.084
	(0.151)	(0.191)	(0.202)	(0.170)	(0.170)	(0.181)	(0.189)	(0.222)
Teen birth four years ago?	0.286[a]	−0.011	0.022	0.209	0.147	0.158	−0.024	−0.057
	(0.148)	(0.193)	(0.206)	(0.168)	(0.169)	(0.170)	(0.202)	(0.231)
Teen births five or more years ago	0.044	0.179	0.178	−0.033	−0.109	−0.085	0.194	0.108
	(0.073)	(0.140)	(0.144)	(0.109)	(0.110)	(0.110)	(0.158)	(0.217)
Covariance parameters								
σ_{μ}^2	–	–	–	0.041	0.083[b]	0.071[a]	0.107	0.296
				(0.039)	(0.042)	(0.043)	(0.210)	(0.272)
$\varrho(1)$	–	–	–	–	−0.167[c]	−0.147[b]	–	–
					(0.063)	(0.071)		
$\varrho(2)$	–	–	–	–	–	0.026	–	–
						(0.080)		
Simulated log likelihood	−1313.43	−1306.85	−1304.13	−1312.67	−1309.83	−1309.75	−1306.27	−1301.09

Note: Estimates based on 384 individual (4118 person-year) observations from the 1979 – 1992 panels of the NLSY. Coefficients for number of siblings, non-intact family at age 14, availability of reading materials at age 14, mother's education (three variables), age, age squared, local birth rate, availability of Ob-Gyn physicians, restrictions on Medicaid funding of abortions, urban residence, local unemployment rate, average annual retail earnings, and an intercept are estimated but not reported. Standard errors appear in parentheses.

[a] Significant at 0.10 level.
[b] Significant at 0.05 level.
[c] Significant at 0.01 level.

cross-section ML probit methods. These specifications, which omit controls for serial correlation, should be helpful in replicating the findings of earlier studies.

The specifications labeled (a) regress current birth outcomes on the full set of independent variables and on the five lagged dependent variables describing the women's early childbearing.[10] The coefficients on these lagged variables provide estimates of the gross differences in fertility behavior between teen mothers and all other women. Consistent with the findings of most previous studies, the estimates indicate that some strong and significant positive associations exist between early childbearing and subsequent fertility for all three racial/ethnic groups. Specifically, white/other women experience significantly higher rates of fertility in each of the first four years following a teen birth but significantly lower rates of fertility thereafter. For black women, there is a very weak and insignificant negative association between teen births and fertility in the following year, a significantly positive association with fertility two years later, and a weak and insignificant association thereafter. For Hispanic women, significantly positive associations are estimated for fertility in the first, second, and fourth years following a teen birth.

The next specifications (b) in each table add general measures for lagged births to the measures for early births. This permits an examination of the subsequent childbearing patterns of teen mothers relative to those of young adult mothers. For all three racial/ethnic groups, the added general lag variables are jointly significant with a similar coefficient pattern — a negative association in the first year after a birth, positive coefficients in the second through fourth years, and a negative association thereafter. The results from specification (b) also indicate that white/other and Hispanic teen mothers have higher rates of subsequent fertility than either women who delay their births until after age 20 or other women generally and that teen births among these two groups are more closely spaced than young adult births. In particular, the first and fourth lags in teen births are significantly positive and the other lag coefficients are weakly positive for white/other women, while the first and second teen fertility lags are significantly positive for Hispanic women. A different relative picture appears for black women in specification (b) where the subsequent fertility behavior of teenage mothers is statistically indistinguishable from that of young adult mothers.

Specification (c) in Tables 1— 3 includes controls for lagged births generally as well as lagged higher parity and teenage births. For all three racial/ethnic groups, the added coefficients are again jointly significant. For all three groups, the significant coefficients from the general fertility lags are positive while the significant coefficients for the higher parity birth lags are negative. The results are consistent with a "two-child norm" and indicate that while subsequent childbearing is higher among women who have already become mothers than other women, childbearing otherwise decreases with parity. For white/other and Hispanic women, the differential effects of teen births are weaker than those estimated in specification (b). For both groups of women, the coefficient on the first teen birth lag is significantly positive, though smaller than in the previous specification, and the coefficients on the remaining lags are insignificant. For black women, the coefficient on the three-year lag in teen births becomes significantly negative while the other coefficients remain insignificant.

To summarize the results so far, estimates from models that omit controls for serial correlation indicate that teen mothers of all three racial/ethnic groups go on to bear more children overall than other women. White/other and Hispanic teen mothers also appear to have more children and more closely spaced births than women who become mothers after age 20. For black women, teen mothers appear to experience slightly lower rates of subsequent fertility than older mothers.

The remaining columns in Tables 1— 3 report estimates from respecified versions of the previous models that incorporate alternative controls for serial correlation. The specifications have been estimated using a two-stage (relatively efficient) application of the MSM panel probit procedure (see Appendix A for details).

The first of these specifications (d) takes the basic model with controls for early childbearing but no other fertility measures and reparameterizes the error term to incorporate a permanent random effect (respecifies the error term for woman *i* in year *t* as $\varepsilon_{i,t} = \mu_i + v_{i,t}\ where\ E(\mu_i) = E(V_{i,t}) = 0, Var(\mu_i) = \sigma_\mu^2, Var(v_{i,t}) = 1 - \sigma_\mu^2, and\ Cov(\mu_i, v_{i,t}) = 0$). The random effect controls for unobserved and imperfectly measured time-invariant individual characteristics such as overall fecundity, preferences for children, and labor market skills. When the results from specification (d) are compared with those from (a), the null hypothesis of no random effects is clearly rejected for all three groups of women. The respecification also leads to noticeable changes in the estimated relationship between early and later fertility with the coefficients on lagged teen fertility becoming more negative for all three racial/ethnic groups. For white/other women, only the third lag on teen births remains significantly positive, and the coefficient on teen births lagged five or more years becomes more strongly negative. For black women, the coefficient on the final lag term also becomes significantly negative.

The next column reports results from a respecification of model (d) in which the transitory component of the error term is parameterized to follow an autoregressive process of order one (an AR(1) process). This relaxes the equicorrelation assumption of the previous specification by permitting decay over time in the effects of some unobserved variables. The specification may capture the effects of biological processes such as the period of subfecundity associated with breast-feeding or gradual changes in economic and social opportunities. For all three groups of women, likelihood ratio tests and estimates of significantly negative AR(1) correlation coefficients lead to a rejection of the equicorrelation assumption. The estimated effects of early fertility appear to be sensitive to the particular method used to control for serial correlation. Specifically, the coefficient for the first lag in teen births becomes more positive for all three groups with the coefficients for white/other and Hispanic women becoming significant. The coefficients for the remaining lags in teen births become more negative.

In specification (f), the transitory component of the error term is reparameterized to follow an AR (2) process. The added correlation coefficient is positive and statistically significant for white/other and black women but weak and insignificant for Hispanic women. Despite the significant AR (2) coefficients, there are few substantive changes in the estimated effects of early fertility for any of the groups. For white/other women, the coefficients on the first two lags become slightly more negative while the remaining coefficients become more positive; for black and Hispanic women, a handful of marginally significant coefficients from the previous specification lose their significance.

The final specifications in Tables 1 — 3 control for the women's observed fertility histories as well as for serial correlation in the unobserved variables. Specification (g) includes a permanent random effect and lags for all births (i.e., adds a random effect to model (b)). The variance coefficient for the permanent component is significant for white/other and black women but insignificant for Hispanic women. Including the component dramatically changes the coefficients on lagged births for white/other and black women with all ten coefficients in each specification becoming negative and most becoming significantly so.

Specification (h), which adds lags in higher parity births to the previous model, nests specifications (a) through (d) as well as (g) as special cases. The coefficients on the added variables are individually insignificant in all three tables, and the inclusion of these variables does not significantly improve the likelihood relative to (g). For all three groups, however, the likelihood is improved relative to models (a) through (d). The coefficient for the variance of the random effect is significant only for white/other women. Across all groups, the coefficients for early births are mostly similar in sign and magnitude to the estimates from model (g), although a few coefficients change their significance levels.

A final specification which adds an AR(1) transitory component to model (g) is estimated for white/other women.[11] While the random effects variance coefficient remains significant, the AR(1) correlation coefficient is not. Relative to model (g), the estimated effects of early births for white/other women decrease in magnitude with the coefficients on two previously significant lag terms losing their significance. For the general lag variables, the coefficients mostly increase in magnitude relative to (g), and all remain significantly negative.

The results from the preceding tables are reported in terms of coefficients on distributed lags and, in some cases, lags interacted with other variables. As such, direct interpretation of the results is difficult. To illustrate better the implications of alternative models, Table 4 lists age-specific predictions of birth probabilities and achieved parities for women who experienced first births at ages 17 and 20. By comparing outcomes across these two groups of women, we can see what various models tell us about the subsequent fertility effects of postponing a teen birth by a few years.[12]

The first results in Table 4 are produced using specification (c), which contains extensive controls for fertility history but no controls for serial correlation. For white/other women, the predicted profiles indicate that women who delay their first birth until age 20 have higher annual fertility rates than the teen mothers from age 22 on. At each reported age, the teen mothers have higher achieved parities with the gap narrowing from 0.487

children at age 21 to 0.160 children at age 28. Although black women have higher overall levels of fertility on average than white/other women, the estimates from specification (c) lead to essentially the same relative outcomes — mostly higher age-specific fertility rates for young adult mothers but higher, albeit gradually narrowing, achieved parities for teen mothers. The predictions for Hispanic women, however, are very different. In particular, the annual fertility rates for Hispanic teen mothers are higher than those for young adult mothers at every age except 22 — 24. The difference in achieved parities at age 28 (0.592) is little different than the difference at age 21 (0.648). Overall, the results for all three groups are consistent with the early literature's findings of higher achieved parities for teen mothers.

The next set of results are based on specification (h) which includes controls for both fertility history and serial correlation. Relative to the previous results, these specifications predict mostly lower annual fertility rates. However, the decreases are generally sharper for the younger than the older mothers. Indeed, for white/other and black women, the specifications predict much lower rates of

Table 4. Predicted subsequent fertility of women with first births at ages 17 and 20

Specification	Outcome	Age at first birth	Age										
			18	19	20	21	22	23	24	25	26	27	28
Non-black, non-Hispanic (white/other) women													
Model (c)	Birth probabilities	17	0.092	0.136	0.172	0.163	0.076	0.084	0.087	0.089	0.093	0.096	0.099
		20	–	–	–	0.077	0.214	0.178	0.155	0.096	0.101	0.103	0.103
	Number of children	17	1.092	1.228	1.400	1.564	1.640	1.723	1.811	1.899	1.992	2.088	2.188
		20	–	–	1.000	1.077	1.291	1.469	1.625	1.721	1.822	1.925	2.028
Model (h)	Birth probabilities	17	0.018	0.032	0.052	0.059	0.031	0.041	0.052	0.062	0.073	0.083	0.091
		20	–	–	–	0.075	0.151	0.150	0.138	0.102	0.113	0.120	0.124
	Number of children	17	1.018	1.050	1.101	1.161	1.192	1.233	1.285	1.347	1.420	1.503	1.594
		20	–	–	1.000	1.075	1.225	1.376	1.514	1.616	1.729	1.849	1.973
Black women													
Model (c)	Birth probabilities	17	0.100	0.230	0.133	0.143	0.170	0.160	0.148	0.132	0.113	0.092	0.072
		20	–	–	–	0.092	0.244	0.191	0.146	0.145	0.121	0.099	0.077
	Number of children	17	1.100	1.330	1.463	1.605	1.775	1.935	2.083	2.215	2.328	2.420	2.492
		20	–	–	1.000	1.092	1.336	1.527	1.673	1.818	1.939	2.038	2.115
Model (h)	Birth probabilities	17	0.040	0.109	0.076	0.086	0.101	0.106	0.107	0.104	0.097	0.086	0.074
		20	–	–	–	0.117	0.245	0.206	0.167	0.161	0.145	0.126	0.106
	Number of children	17	1.040	1.149	1.225	1.311	1.412	1.518	1.626	1.730	1.827	1.913	1.987
		20	–	–	1.000	1.117	1.362	1.568	1.735	1.896	2.041	2.167	2.273

Table 4 (continued)

Specification	Outcome	Age at first birth	Age										
			18	19	20	21	22	23	24	25	26	27	28
Hispanic women													
Model (c)	Birth probabilities	17	0.135	0.247	0.161	0.200	0.145	0.154	0.133	0.121	0.099	0.081	0.062
		20	–	–	–	0.096	0.202	0.169	0.187	0.090	0.085	0.066	0.051
	Number of children	17	1.135	1.382	1.544	1.744	1.889	2.043	2.176	2.297	2.396	2.477	2.539
		20	–	–	1.000	1.096	1.298	1.466	1.654	1.744	1.829	1.896	1.947
Model (h)	Birth probabilities	17	0.067	0.142	0.100	0.130	0.095	0.101	0.096	0.091	0.080	0.069	0.057
		20	–	–	–	0.081	0.173	0.149	0.163	0.081	0.077	0.065	0.054
	Number of children	17	1.067	1.209	1.309	1.439	1.534	1.635	1.731	1.822	1.901	1.970	2.027
		20	–	–	1.000	1.081	1.254	1.403	1.565	1.646	1.723	1.788	1.842

Note: Estimates based on 1979 – 1992 data from the NLSY. Predictions are evaluated the race-specific means of all of the independent explanatory variables other than fertility history and age (the variable means are listed in Appendix B). Predictions at each age in incorporate the possibility of births at prior ages.

fertility among the teen mothers than the young adult mothers. The differences are such that by age 22 the predicted cumulative fertility of the age-20 white/other mothers actually exceeds that of the age-17 mothers; the corresponding cross-over point for black women is age 23. For Hispanic women, the achieved parities of younger and older mothers do not cross-over, though the remaining differences are narrower than in the previous specification. [13]

5. Conclusions

This paper estimates panel probit models of annual birth outcomes and examines the relationship between teenage and young adult childbearing. Using models which omit controls for women's fertility histories and assume that the unobserved determinants of fertility are serially uncorrelated, the paper replicates existing findings of positive effects of early births on later childbearing. In subsequent models, however, it finds strong evidence of serial correlation as well as evidence that the estimated relationship between teen and young adult childbearing is sensitive to controls for serial correlation and other fertility behavior. For white and black women, adding these controls leads to significant negative estimates of the effect of early fertility on later childbearing and birth spacing. In the case of Hispanic women, the controls lead to estimates which are consistent with either positive or negative effects of teen childbearing. Overall, there is little evidence that postponing births from just before to just after age 20 reduces subsequent childbearing among women who are otherwise disposed to become mothers.

Like earlier studies by Heckman et al. (1985) and Heckman and Walker (1991), the paper demonstrates the importance of controlling for unobserved heterogeneity and distinguishing between particular sources of heterogeneity in models of fertility. As with several other recent studies, it also shows that the estimated consequences of teen childbearing depend critically upon assumptions regarding such heterogeneity. In terms of policy, the results caution us that programs which attempt to reduce adolescent fertility without addressing its underlying causes may only marginally affect women's subsequent economic and social well-being.

Endnotes

[1] Examples of these recent studies include Bronars and Grogger (1994), Geronimus and Korenman (1992), Hoffman et al. (1993), Hotz et al. (1995), Klepinger et al. (1995 a, b), Olsen and Farkas (1989), and Ribar (1994, 1995).

[2] Examples of theoretical life-cycle fertility models which incorporate timing effects include Blackburn et al. (1993), Happel et al. (1984), Moffitt (1984a), and Vijverberg (1984); see also the review of these models by Montgomery and Trussell (1986).

[3] Differing estimates of the human capital consequences of early fertility have been reported by Bronars and Grogger (1994), Geronimus and Korenman (1992), Hofferth (1987), Hoffman et al. (1993), Klepinger et al. (1995 a, b), Olsen and Farkas (1989), Ribar (1994, 1995), and Upchurch and McCarthy (1990).

[4] Moffitt (1984b) estimated panel probit models of married women's annual birth outcomes. His study did not explicitly examine the effects of birth timing on subsequent fertility. It did, however, control for and find evidence of serial correlation. It also reported "a strong education-timing interaction, showing that more educated women begin child-bearing later and have children at a faster rate once they begin".

[5] If the source of serial correlation can be expressed in terms of a factor analytic structure (e.g., as in the single-factor random effects model), evaluation of the integral is greatly simplified and can be accomplished using standard numerical quadrature methods (for the random effects example, see Butler and Moffitt 1982).

[6] In the initial survey, blacks, Hispanics, disadvantaged white youth, and military personnel were over-sampled. Weights (not used here) are available to make the data nationally representative. For each household sampled in 1979, the NLSY gathered data for all individuals aged 14 — 21 within the household. The present analysis does not account for possible household-specific correlations across individuals (these correlations are extensively examined by Ribar 1995).

[7] The restriction eliminates data for 1 woman who had a birth in the calendar year she turned 14 and 13 women who had births in the calendar year they turned 15. The exceedingly small sample probabilities for these births made the events unsuitable for analysis.

[8] This restricts births lagged five or more years to have identical effects on the dependent variable. The lag restriction was selected after some experimentation. Alternative lag specifications do not substantially change the study's findings. Because women with births prior to age 15 have been removed from the analysis, lags describing fertility before this age have been initialized at zero.
[9] Computation time for the MSM procedure rises nearly proportionately with the number of model coefficients. The study's design, which involves estimation and comparison of results from numerous MSM specifications, necessitates that a relatively parsimonious set of independent variables be used. Several other independent controls — e.g., variables for father's education, detailed family structure at age 14, religion and religiousness, other local economic conditions, and access to other reproductive health services — have been included in various specifications without substantially altering the results.
[10] For brevity, coefficients for the independent variables are not reported. These results are available from the author upon request.
[11] Imprecision in (weak identification of) the covariance parameters led to very slow convergence in the MSM procedure for white/other women. Similar models failed to converge for black or Hispanic women. Accurate simulation on the likelihood was hampered by the high degree of serial correlation. Despite the fact that model (i) nests model (g), the simulated likelihood for (g) is slightly higher than for (i). Note that the likelihood requires simulations of the full sequence probabilities for each individual; the problems in this calculation are not transmitted to the MSM estimates which rely on independently simulated and smaller dimension transition probabilities. The results for specification (i) are reported as a general robustness check.
[12] The predictions in Table 4, which are computed using point estimates of the model coefficients and not accompanied by confidence intervals, should be interpreted with caution.
[13] To conserve space, predictions from other specifications are not reported. The implications of the results from specifications (d), (e), and (1) are relatively clear. For specification (g), the predictions were qualitatively similar to those from model (h).

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Appendix A

Description of method of simulated moments panel probit estimator

Consider the panel probit model

$$y^*_{i,t} = x_{i,t}\beta + \varepsilon_{i,t}$$
$$y_{i,t} = \begin{cases} 1 & \text{if} \quad y^*_{i,t} \leq 0 \\ 0 & \text{otherwise} \end{cases}, \qquad \text{(A.1)}$$

where $i = 1, N$, $t = 1, T$, $x_{i,t}$ denotes a vector of observed individual- and possibly period-specific independent variables, and $\varepsilon_{i,t}$ denotes an individual- and period-specific error term independently and identically distributed across i such that $[\varepsilon_{i,1} \ldots, \varepsilon_{i,T}] = \varepsilon_i \sim N(0, \Sigma)$. For simplicity, the model abstracts from lags in the outcome variable $y_{i,t}$. Let θ be vector which consists of the elements of β and Σ, $S_{i,t} = \{y_{i,t}, \ldots, y_{i,t}\}$i be the set of choices observed for i through period t, and $\text{Prob}(S_{i,t}|X_i, \theta)$ be the probability of observing that set of choices.

Keane (1992) noted that the log likelihood for this model could be expressed as a sum of a series of transition probabilities such that

$$LLF(\hat{\theta}) = \sum_{i=1}^{N} \ln \text{Prob}\,(S_{i,T} \mid X_i, \hat{\theta})$$

$$= \sum_{i=1}^{N} \sum_{t=1}^{T} \{y_{i,t} \ln \text{Prob}\,(y_{i,t} = 1 \mid S_{i,t-1}, X_i, \hat{\theta}) \qquad \text{(A.2)}$$

$$+ (1 - y_{i,t}) \ln \text{Prob}\,(y_{i,t} = 0 \mid S_{i,t-1}, X_i, \hat{\theta})\}$$

with the associated score

$$\frac{\partial LLF(\hat{\theta})}{\partial \hat{\theta}} = \sum_{i=1}^{N} \sum_{t=1}^{T} \{W^1_{i,t}[y_{i,t} - \text{Prob}\,(y_{i,t} = 1 \mid S_{i,t-1}, X_i, \hat{\theta})]$$

$$+ W^0_{i,t}[(1 - y_{i,t}) - \text{Prob}\,(y_{i,t} = 0 \mid S_{i,t-1}, X_i, \hat{\theta})]\} \; , \qquad \text{(A.3)}$$

where

$$W^m_{i,t} = \frac{\partial \text{Prob}\,(y_{i,t} = m \mid S_{i,t-1}, X_i, \hat{\theta}) / \partial \hat{\theta}}{\text{Prob}\,(y_{i,t} = m \mid S_{i,t-1}, X_i, \hat{\theta})} \quad \text{for} \quad m = 0\;, \quad 1\;. \qquad \text{(A.4)}$$

The score (A.3) can be used to form a conventional method of moments estimator for $\hat{\theta}$. Given a consistent $\hat{\theta}$, (A.4) describes the optimal weights for this estimator. To construct an MSM estimator (McFadden 1989), Keane replaced the transition probabilities in (A.3) with simulations from a highly accurate, recursive algorithm developed by Geweke, Hajivassiliou, and Keane (see Geweke et al. 1994; Hajivassiliou et al. 1994; Keane 1992, 1994) and used independent simulations of the transition probabilities to construct estimates of the optimal weights (A.4). Keane (1994) showed that this estimator preserves the desirable properties of McFadden's MSM multinomial probit estimator — namely, that for a fixed simulation size, the estimator is consistent and asymptotically normal in *N*. Keane also reported Monte Carlo evidence indicating that the procedure worked well for relatively small simulations sizes (10 draws per simulated probability) even when applied to models with high degrees of serial correlation.

This study uses an iterative MSM estimation procedure written by the author in FORTRAN. The program contains a GHK binary probit simulation routine patterned after a more general Gauss procedure written by Geweke et al. (1994). Following McFadden's (1989) recommended estimation strategy, the MSM program was applied in two stages. In the first stage, weights were simulated using best-guess starting values for $\hat{\theta}$ and a simulation size of 10 draws per probability. Consistent, but inefficient, estimates of $\hat{\theta}$ based on 10 draws were then obtained from the MSM program. For the second stage, the program was re-run using a simulation size of 20 and optimal weights generated with the first-stage estimates of 0. The study reports parameter estimates and estimated standard errors from this second stage. Estimation times at each stage on a small workstation were reasonable ranging from roughly 10 min for first-stage specifications with the fewest parameters and observations up to nearly 2 h for second-stage specifications with the most parameters and observations.

Appendix B

Descriptive statistics for analysis data set

Variable	White/ other women		Black women		Hispanic women	
	Mean	(SD)	Mean	(SD)	Mean	(SD)
Birth	0.082	(0.274)	0.115	(0.319)	0.102	(0.303)
Teen mother	0.169	(0.375)	0.297	(0.457)	0.265	(0.441)
Number of siblings	3.242	(2.105)	4.458	(3.009)	4.248	(2.825)
Non-intact family at age 14	0.287	(0.452)	0.569	(0.495)	0.300	(0.458)
Magazines in home at age 14	0.680	(0.467)	0.399	(0.490)	0.425	(0.494)
Mother's education missing	0.043	(0.203)	0.072	(0.259)	0.055	(0.227)
Mother's years of education	10.974	(3.373)	9.980	(3.750)	7.668	(4.148)
Mother's years of post-secondary education	0.501	(1.271)	0.321	(1.023)	0.188	(0.822)
Birth rate in county of residence	66.096	(11.444)	68.985	(10.405)	75.980	(13.466)
Ob-gyn physicians per 1000 women in county of residence	0.427	(0.293)	0.607	(0.360)	0.507	(0.244)
State restricts Medicaid funding for abortions	0.548	(0.482)	0.666	(0.452)	0.469	(0.487)
Urban residence	0.700	(0.458)	0.803	(0.398)	0.944	(0.230)
County unemployment rate	8.018	(3.387)	7.395	(2.865)	7.936	(3.288)
Annual average retail earnings in county of residence	10.845	(1.728)	11.223	(1.458)	11.845	(1.680)
Observations at						
Age 15	1128		531		384	
Age 16	1128		531		384	
Age 17	1109		523		376	
Age 18	1073		508		365	
Age 19	1021		484		348	
Age 20	976		469		329	
Age 21	924		456		315	
Age 22	884		431		296	
Age 23	833		415		279	
Age 24	807		404		268	
Age 25	738		395		256	
Age 26	664		380		248	
Age 27	427		264		183	
Age 28	221		140		87	
Total person-year observations	11933		5931		4118	

Note: Data from 1979 – 1992 panels of NLSY