A Joint Dynamic Model of Fertility and Work of Married Women

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This article estimates a dynamic model of fertility and labor supply of married women drawn from the National Longitudinal Survey of Young Women, 1968–91. It distinguishes part-time and full-time employment sectors, which differ by pecuniary and nonpecuniary returns and transferability of human capital. The model with unobserved heterogeneity in earning ability and preferences for children fits the data and produces reasonable forecasts of labor force participation in decisions. The estimates unpack important features of the persistence in labor market decisions, intertemporal substitution of leisure over the life cycle, and the effect of work interruptions, due to childbirth, on lifetime utility.

I. Introduction

This article presents and estimates a joint dynamic model of labor supply and fertility of married women. The model assumes that these women behave as though they were solving a finite-horizon, discrete-choice, dynamic programming problem under uncertainty. The structural parameters of the model are estimated using data on a sample of married women from the National Longitudinal Survey (NLS) of Young Women, 1968–91. This study contributes to both the specific literature on female labor force decisions and the general literature on structural modeling of married

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women's behavior (see Blundell and MaCurdy 1999). Its contribution to the specific literature is to distinguish part-time and full-time labor force involvement. In each discrete period, women are assumed to make sequential decisions among three employment states (nonwork, part-time work, and full-time work), which differ by pecuniary and nonpecuniary returns and the extent of transferability of sector-specific human capital. This article's contribution to the general literature is the joint modeling of labor force participation and fertility decisions in a dynamic and stochastic framework. That is, the employment choice is made in conjunction with the decision of whether or not to have a child in each discrete period. Thus, women are faced with six alternative participation-fertility states.

The approach followed in this article relies on economic theory to guide the empirical work and to make predictions. An important benefit of this approach relative to others is that one can interpret all the estimated parameters quite easily because they are directly cast into an economic model of fertility and labor supply. A second benefit is that one can forecast the fertility and labor supply behavior of women given any change in the state of the world through a change in their constraints, even changes along dimensions that are invariant in the data (Wolpin 1996). However, one of the main costs of a structural estimation approach is computational burden. The full-solution approach to structural estimation adopted here requires that the optimization problem be solved numerically for the optimal decision rules of all women in the model, that is, it requires repeated solution of the optimization problem at each trial parameter value. Another problem is the potentially high level of parametrization that may be overly restrictive on behavior.[1]

The structural parameters of the model, which are used to assess the model fit and to perform several simulations, are recovered by repeatedly solving the dynamic program and maximizing a likelihood function that reflects choices and earnings observed for each woman over a variable period (between 8 and 24 years) since marriage.[2] For the years in which earnings are available, only the earnings in the chosen employment sector are known. Because we solve a dynamic program in which part-time and full-time earnings are stochastic, the wage function parameters represent wages offered by employers. The model is shown to fit the data quite well in several directions and produces reasonable forecasts of future labor force participation decisions.

There are four substantive findings that are worthwhile to stress. First,

[1] A number of recent studies have illustrated alternative estimation strategies (Angrist 1995; Keane and Wolpin 1997*b*; Angrist and Krueger 1999).

[2] See Eckstein and Wolpin (1989*b*) and Rust (1994) for surveys of recent developments in the literature concerned with the methods for solving and estimating dynamic structural models.

the observed persistence in full-time employment (which is a common feature of several micro data sets and has been documented by a number of previous studies) is primarily due to large positive full-time and schooling effects on full-time earnings. The relatively lower persistence in part-time employment is due to the small own-experience effect on part-time earnings. For our sample of married women, therefore, labor market transitions seem to be driven by labor market returns rather than work preferences. Second, the wage-experience profile of full-timers is substantially different from that of part-timers. In particular, full-time employment is only mildly substitutable with part-time employment, and leisure appears to be more easily intertemporally substitutable with full-time work than with part-time work. Third, there is a clear relationship between earnings ability and preference for children. Women with a comparative advantage in market work (i.e., with the highest earnings profiles) are those with the lowest marginal utility of children; conversely, women with the strongest preference for children are those with lower earnings profiles. Fourth, part-time employment does not appear to cushion women who were previously working full-time from losses associated with childbearing. In fact, any part-time experience accumulated during the childbearing/childrearing period tends to depress the earnings profiles of mothers who return to full-time employment. Short interruptions of full-time work turn out to be more beneficial (for the mother) around childbirth.

The rest of the article is organized as follows. Section II outlines the background against which the contribution of this article must be seen. Section III presents the behavioral model. Section IV illustrates the solution and the main estimation issues. Section V describes the data and addresses the problem of sample selection that is relevant for virtually all discrete-choice structural dynamic models. Section VI presents the estimation results and the model's ability to fit the data. Section VII discusses the implications of the model in two important areas: (i) the effects of changing women's socioeconomic environment on their employment and fertility patterns and (ii) the efficacy of part-time work as an alternative to work interruptions after childbirth. Section VIII concludes.

II. Background

The theoretical framework of the model that will be developed in Section III combines the basic features of models that describe the life-cycle capital accumulation process with an endogenous labor supply (Weiss 1972; Heckman 1976; Weiss and Gronau 1981) and the basic features of models that emphasize the importance of variations in the value of time to explain individual fertility behavior (Becker 1965; Willis 1973; Heckman and Willis 1976; Hotz 1980). The former class of models argues that labor market participation affects future wages, which then, in turn, affect future participation.

Because the investment return to current work is taken into account in a forward-looking optimizing model, work experience (or accumulated participation) is endogenous. The first empirical study to have incorporated endogenous experience in a model of female labor force participation is that by Eckstein and Wolpin (1989*a*).[3] We build on their contribution by recognizing the importance of hours of work decisions (Heckman and MaCurdy 1980; Blank 1988). Women show more variation than men in their labor supply, not only because many women choose not to be in the labor force, but also because many choose to work part-time (or part-year, or both). With the distinction between part-time and full-time work, we add one salient dimension to the analysis of the sensitivity of women's labor supply to intertemporal variation in wages. The intertemporal substitution in leisure over the life cycle operates through both preferences for and opportunities in the two forms of employment. In addition, the distinction between part-time and full-time work embodies the notion of different degrees of transferability of human capital accumulated in part-time and full-time jobs (Corcoran, Duncan, and Ponza 1983). Women's choices among these employment "sectors" take into account future wage and job opportunities that, in turn, depend on endogenously accumulated sector-specific work experience, as well as on the number of children and on women's ability. Despite the substantial proportion of women working part-time and an ongoing public discussion about the problems and advantages of part-time jobs (Blank 1990; Blossfeld and Hakim 1997),[4] labor supply research that has focused on part-time work is surprisingly scarce. A few studies have investigated part-time work choices with cross-sectional information, including Jones and Long (1979) and Fraker and Moffitt (1988) for the United States, Nakamura and Nakamura (1983) for Canada, and Ermisch and Wright (1993) for Britain. Others have investigated the role of part-time work in a more dynamic context, including Corcoran et al. (1983) and Blank (1989, 1994) for the United States, and Burchell, Dale, and Joshi (1997) for Britain. None of these studies, however, model labor supply and fertility choices as joint decisions.

The latter class of models that underpins our study argues that the

[3] Only a few empirical studies have allowed for work experience to be accumulated endogenously. In the female labor supply literature, see Van der Klaauw (1996) and Altug and Miller (1998). Exceptions in the literature on male labor supply and career decisions include Shaw (1989), Wolpin (1992), Keane and Wolpin (1997*a*), and Sauer (1998). For a comprehensive survey of structural dynamic models of labor supply, see Blundell and MaCurdy (1999, sec. 8).

[4] Over the period 1968–92, between 25% and 30% of employed women have worked part-time according to the official definition of part-time work, i.e., less than 35 hours per week (Blank 1994; Drobnic and Wittig 1997).

fertility decision can be viewed as an economic one, with the cost of having a child being measured by the mother's forgone (present and future) earnings and the parents deriving pure consumption benefits from children.[5] While the joint nature of fertility and labor supply decisions has long been recognized in both static models (Cain and Dooley 1976; Schultz 1980) and dynamic models (Moffitt 1984; Hotz and Miller 1988), there is no empirical study that has examined the joint dynamic relationship between fertility and labor supply with endogenously accumulated work experience. This is important, because sector-specific participation at one point in time increases the stock of sector-specific human capital, which, in turn, affects the potential wage in both employment sectors at all later points in time and thus affects the cost of having a child. Neither is there any structural forward-looking optimizing model of fertility and labor supply. This is important for at least three reasons. First, because the estimation framework imposes several restrictions on the model, we can test whether such a restricted model fits the observed data patterns adequately with a parsimonious set of information. Second, because we specify a forward-looking model, dynamics play a role through expectations and not only through past behavior or past wage realizations. For example, our model allows us to study how much and in what way current fertility choices respond to changes in sector-specific (future) wage returns to work experience. Third, because we explicitly solve an optimization problem, we can quantify the behavioral effect of changing specific parameters of interest. For example, as fertility and work are interrelated, we can estimate the effect of an intervention that alters the sector-specific wage profiles, such as a labor income tax/subsidy or a contractual pay change, on subsequent fertility choice decisions. Previous research that has treated fertility and participation decisions in isolation has been limited in addressing such issues.[6]

[5] For a review, see Lehrer, Grossbard-Shechtman, and Leasure's (1996) comment on Friedman, Hechter, and Kanazawa's (1994) theory of the value of children and the references cited therein.

[6] Angrist and Evans (1998) use a different methodological approach to investigate the effect of fertility on labor supply. To this purpose, they apply an instrumental variables strategy based on the sibling sex composition in families with two or more children. Notice that their study does not focus on the causal link running from labor supply to fertility, while the joint determination of fertility and work decisions is one of the primary motivations of our research. Moreover, estimates of the effect of going from two to more than two children do not necessarily generalize (according to Angrist and Evans [1998] in 1980 only 40% of women with two children had a third child, and only 40% of women aged 21–35 had two or more children). Any generalization requires that we learn about the mechanism generating the mother's (or parents') response to the fertility change at each initial parity level. See also Rosenzweig and Wolpin (2000).

III. The Model

This section presents a dynamic stochastic model of female labor force participation and fertility. The model is restricted to married women.[7] Each woman has a decision horizon that begins at age A_0, the age at which she gets married, and ends at age 65. At each age t, a woman chooses whether she works part-time ($p_t = 1$) or not, whether she works full-time ($f_t = 1$) or not, and whether she has a child ($n_t = 1$) or not. A birth may occur in any period during the fertile stage, which exogenously ends at age 40 for the wife. Therefore, a woman faces six mutually exclusive alternatives denoted by j: $j = 1$ if no birth and nonwork; $j = 2$ if birth and nonwork; $j = 3$ if no birth and part-time work; $j = 4$ if birth and part-time work; $j = 5$ if no birth and full-time work; and $j = 6$ if birth and full-time work. Let $d_{jt} = 1$ if alternative j is chosen at time t, and $d_{jt} = 0$ otherwise. All alternatives are mutually exclusive, implying $\sum_{j=1}^{6} d_{jt} = 1$. At any age t, the objective of the wife is to maximize the expected present value of remaining lifetime utility,

$$E_t\left[\sum_{t=A_0}^{65} \delta^{t-A_0}\, U_t\Big(c_t, p_t, f_t, n_t, X_{t-1}^{p}, X_{t-1}^{f}, N_t, S\Big)\right], \tag{1}$$

with respect to p_t, f_t, and n_t for ages $t = A_0, \ldots, 40$ and with respect to p_t and f_t for ages $t = 41, \ldots, 65$. The variables are defined as follows: c_t is the level of goods consumption at t; p_t is a dichotomous variable equal to unity if the woman works part-time and equal to zero otherwise; f_t is a dichotomous variable equal to unity if the woman works full-time and equal to zero otherwise; $n_t = 1$ indicates a birth at t and $n_t = 0$ indicates no birth; X_{t-1}^{p} and X_{t-1}^{f} denote the number of prior periods the woman has worked part- and full-time, respectively; N_t indicates the total number of children in the household at age t; S is the completed level of schooling of the wife; δ is the subjective discount factor; and $E[\cdot]$ is the mathematical expectations operator.

Without saving or borrowing, household choices are resource-constrained each period by[8]

$$y_t^{h} + y_t^{p} p_t + y_t^{f} f_t = c_t + b_1 p_t + b_2 f_t + b_3 N_t, \tag{2}$$

[7] Allowing the model to include other family types creates major complications both in solving the dynamic program and in estimation, because it requires to treat the marital decision as endogenous. This leads, however, to the important issue of sample selection. We shall return to this point in Sec. V.

[8] Women's work behavior around childbirth may be related to savings accumulation and household wealth. The assumption that the household does not carry over any debt incurred during a period to the next period is extreme. In fact, as noted by Eckstein and Wolpin (1989*a*), if U_t is linear and additive in consumption, the problem becomes a problem of wealth optimization modified

where y_t^h is husband's earnings; y_t^p and y_t^f denote wife's earnings in part- and full-time employment, respectively; b_1 and b_2 represent the corresponding fixed costs of work; and b_3 is the goods cost per child.[9]

Because the size of the choice set is directly related to the dimension of the integrations performed to solve the dynamic program—which is what Bellman (1957) called the "curse of dimensionality"—we assume that the husband's labor income is determined by (some of) his wife's observable characteristics.[10] To forecast the expected value of husband's earnings at a specific age t, each woman forms her own expectations about male earnings potential according to

$$\bar{y}_t^h = Z_t'\phi, \tag{3}$$

where ϕ is a suitable parameter vector, and Z_t is a set of sociodemographic variables of the wife, such as age at marriage, schooling, and race. Equation (3) reflects the hypothesis that women with similar observable characteristics marry husbands with similar expected earnings (Van der Klaauw 1996). Each woman would then form these expectations by evaluating the earnings of husbands of women with choice histories and characteristics that are most similar to hers.

The wife's current earnings are both endogenous and stochastic. The part-time and full-time employment sectors have different wage structures. Standard sector-specific earnings functions are modified to allow for cross-specific experience effects,[11]

$$\log(y_t^r) = a_0^r + a_1^r X_{t-1}^p + a_2^r (X_{t-1}^p)^2 + a_3^r X_{t-1}^f + a_4^r (X_{t-1}^f)^2 + a_5^r S + \epsilon_t^r \tag{4}$$

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by the psychic value of work and children. Moreover, if the monetary costs of children and work are not observed but treated as parameters (as in [2]), they cannot be distinguished from their psychic values, except by functional form. See Sec. IV for a discussion of the identification issue.

[9] Rather than a technologically determined constant cost, it is, of course, more appropriate to assume that the maintenance cost of a child varies with his or her age (Hotz 1980; Hotz and Miller 1988). However, this will expand the dimensionality of the state space for children quite dramatically: from a state space that is equal to the number of children in any period (i.e., T possible states in period T) to a state space in which the size depends on potential sequences of children. For example, if the maintenance cost changes each period of a child's life, we would have 2^T additional possible state values to compute at period T.

[10] Under the assumption that husband's wages are realized only after female participation and fertility decisions are made, and if the form of the utility function in (1) is linear and additive in consumption (as given in [5] below), only the husband's expected earnings, $\bar{y}_t^h$, enter the decision process (see Eckstein and Wolpin 1989*a*; Van der Klaauw 1996). The form of the utility function is discussed below.

[11] Ermisch and Wright (1993) provide several reasons of why wages in part-time jobs are different from wages in full-time jobs. For example, the cost of

for $r = p, f$. The technology shocks, ϵ_t^p and ϵ_t^f, capture random fluctuations in earnings that are independent of the individual decision process. They have zero mean, finite variance, and a nonzero contemporaneous covariance.

As pointed out earlier, if the instantaneous utility function, U_t, is linear and additive in consumption, the problem becomes that of wealth maximization modified by the psychic value of children and work. Per period utility at any age t is then assumed to have the following form:

$$\begin{aligned} U_t = {} & c_t + \alpha_1 p_t + \alpha_2 f_t + (\alpha_3 + \epsilon_t^n)N_t + \alpha_4 N_t^2 \\ & + \beta_1 c_t p_t + \beta_2 c_t f_t + \beta_3 c_t n_t + \beta_4 p_t n_t + \beta_5 f_t n_t \\ & + (\gamma_1 X_{t-1}^p + \gamma_2 X_{t-1}^f + \gamma_3 S + \gamma_4 N_t + \gamma_5 N_t^2)p_t \\ & + (\gamma_6 X_{t-1}^p + \gamma_7 X_{t-1}^f + \gamma_8 S + \gamma_9 N_t + \gamma_{10} N_t^2)f_t. \end{aligned} \tag{5}$$

The utility in (5) is decreasing in p_t and f_t, reflecting disutility of working part- and full-time, and increasing in consumption, c_t. A larger number of children provides a higher instantaneous utility ($\alpha_3 > 0$), at a decreasing rate ($\alpha_4 < 0$). Each couple has costless control over completed fertility, and births are timed without error (Wolpin 1984; Ahn 1995). Note, however, that the marginal utility of children is random through the taste parameter ϵ_t^n. Household consumption interacts with the labor market decisions through β_1 and β_2, and with the birth decision through β_3. The interactions between work and fertility choices are captured by β_4 and β_5. Letting the labor market decisions interact with prior experience (γ_1, γ_2, γ_6, and γ_7) implies that the utility function is not assumed to be intertemporally separable, as long as (at least) one of these parameters differs from zero: a positive value of the parameter(s) may be interpreted as habit formation in the labor market, whereas a negative value may reflect an increasing disutility of working with previous work effort or increasing propensity to substitute nonmarket time in subsequent periods. Prior experience accumulated in part- and full-time jobs is allowed to have differential effects on utility depending on current participation (i.e., γ_1, γ_6 can be different from

recruiting and training may be a significant impediment to hiring part-time workers (Montgomery 1988). In addition, there may be compensating wage differentials, with attributes of part-time work (e.g., shorter working hours that mesh with home responsibilities) being substitutable for wages. A different wage structure may also arise because of spatial constraints in the labor supply of part-time workers. If part-time workers have a more restricted geographic mobility than full-time workers, and employers exercise a degree of monopsony power in the local labor market, then profit maximization entails paying part-time workers a lower wage. Finally, on the demand side, there is scope for substitution between part-timers and full-timers. Firms can alter the employment mix of part-time and full-time workers according to both the supply and demand conditions in the labor and product markets (Ermisch and Wright 1993; Friesen 1997). In both cases, part-timers will rarely be observed earning more than full-timers.

γ_2, γ_7). The parameters γ_3, γ_8 and $\gamma_4, \gamma_5, \gamma_9, \gamma_{10}$ capture the effects that schooling and children, respectively, have on utility through their interactions with current participation.

The decisions made at age t depend on the fertility and employment histories up to that point in time. This history defines the state at which a woman starts a new period. The state at any age t contains all the relevant history of choices that enter the current utility as well as the realizations of the three shocks up to t. We denote the state space by $\Omega_t = (N_{t-1}, X^p_{t-1}, X^f_{t-1}, S, A_0, \bar{y}^h_t, \epsilon^n_t, \epsilon^p_t, \epsilon^f_t)$. Let the value function $V_t(\Omega_t)$ be the maximal expected present value of lifetime utility as in (1) given the woman's state Ω_t. The value function can be written as the maximum over alternative-specific value functions, that is,

$$V_t(\Omega_t) = \max\left[V_{1t}(\Omega_t), \dots, V_{6t}(\Omega_t)\right], \qquad t = A_0, \dots, 40,$$

and

$$V_t(\Omega_t) = \max\left[V_{1t}(\Omega_t), V_{3t}(\Omega_t), V_{6t}(\Omega_t)\right], \qquad t = 41, \dots, 65,$$

where $V_{jt}(\cdot)$ is the value function if the woman chooses alternative j. Each of these alternative-specific value functions obey the Bellman equation (Bellman 1957):

$$\begin{aligned} V_{jt}(\Omega_t) &= U_{jt}(\Omega_t) + \delta E V_{t+1}(\Omega_{t+1} \mid \Omega_t, d_{jt} = 1), \qquad t < 65 \\ V_{j,65}(\Omega_{65}) &= U_{j,65}(\Omega_{65}), \end{aligned} \tag{6}$$

with $j = 1, \dots, 6$ for $t = A_0, \dots, 40$ and $j = 1, 3, 5$ for $t = 41, \dots, 65$. The expectation in (6) is taken over the distribution of the random components of Ω_{t+1} conditional on Ω_t and $d_{jt} = 1$. The time-varying predetermined state variables such as number of children and sector-specific experience evolve independently of the shocks:[12]

$$\begin{aligned} N_t &= N_{t-1} + n_t \\ X^p_t &= X^p_{t-1} + p_t \\ X^f_t &= X^f_{t-1} + f_t. \end{aligned} \tag{7}$$

To close the model, the three shocks $(\epsilon^n_t, \epsilon^p_t, \epsilon^f_t)$, or ε_t, are assumed to be jointly normal, $\mathcal{N}(0, \Sigma)$ and serially uncorrelated. This assumption simplifies the conditional joint density function at $t + 1$, $\varphi(\varepsilon_{t+1} \mid \Omega_t, d_{jt} = 1; \Sigma)$, to $\varphi(\varepsilon_{t+1}; \Sigma)$. Because fertility and participation decisions may be affected by some common unobserved variables, the random components

[12] Note that substituting (2)–(4) and (7) into (5) implies that the alternative-specific per period utility, U_{jt} in (6), becomes a function of the entire state space.

of the model are allowed to be contemporaneously correlated, that is Σ is a full-rank variance-covariance matrix.[13]

IV. Solution and Estimation Issues

The standard solution method for finite horizon dynamic programs is backward recursion. To outline the decision process underlying our solution method, denote the set of predetermined elements of the state space at age t, that is Ω_t net of the random shocks, as Ω_t^d, and consider a woman who begins her marriage at age A_0. Given $\Omega_{A_0}^d$, the woman draws three random shocks from the joint distribution $\varphi(\varepsilon_{A_0}; \Sigma)$, calculates the (six) current utilities and the alternative-specific value functions, and chooses the alternative that yields the highest value. The state space is then updated according to the alternative chosen, and the process is repeated.[14] Exact numerical solution is carried out by backward recursion.

The solution of the optimization program serves as the input into estimating the structural parameters of the model given data on choices and earnings. To see this, note that we have information on the set of choices and accepted wages over time for each woman who is assumed to solve the model described above. Let k_t denote the combination of choices and earnings observed at age t. Thus, $k_t = (d_{jt}, j = 1, 2)$, $k_t = (d_{jt}, y_t^p d_{jt}, j = 3, 4)$, and $k_t = (d_{jt}, y_t^f d_{jt}, j = 5, 6)$. The probability that a woman is observed to choose alternative j with choice-earnings combination k_t at age t is defined by $Pr(k_t \mid \Omega_t^d) = Pr\{\max_j [V_{jt}(\Omega_t)]\}$. Because of the serial independence of the shocks, the probability of any sequence of choices and earnings is then given by

$$Pr(k_{A_0}, \dots, k_T \mid \Omega_{A_0}^d) = \prod_{t=A_0}^{T} Pr(k_t \mid \Omega_t^d). \tag{8}$$

The likelihood function for a sample of I wives is the product of the probabilities in (8) over the I women. The solution to the individual's optimization problem provides the choice probabilities that are in the right-hand side of (8).[15] Estimation is an iterative process that involves, first, solving numerically the dynamic program for given parameter values

[13] The assumption of serial independence implies that the expected value operator in (6) is not time subscripted. Because of the same assumption, past realizations of the shocks are not included in the state space Ω_t.

[14] After her fortieth birthday, the wife draws only the two technology shocks, ϵ^p and ϵ^f, from φ and computes three alternative-specific value functions.

[15] Choice probabilities are constructed by repeatedly drawing three (or two, after age 40 of the woman) current-period technology shocks and using a kernel smoothing function (McFadden 1989; Pakes and Pollard 1989; Rust 1994).

and, then, computing the likelihood function, until the likelihood is maximized.[16]

The likelihood function obtained from (8) applies to a sample of women who differ in their state at the beginning of marriage but are otherwise homogeneous. Clearly, differences at the beginning of marriage may arise because of different labor market and fertility choices prior to marriage as well as innate "ability." The first source of heterogeneity is observable in the data and will be accounted for in $\Omega^d_{A_0}$. To allow for the possibility that women have different "ability," we assume that there are H types of women ($H < I$), with different wage and preference parameters and comprising $\mu_1, \ldots, \mu_H$ proportions of the population, respectively. Thus, individuals may have comparative advantages among the different alternatives, including in raising children and in acquiring market experience, which are known to them. For each type h the contribution to the likelihood function in (8) becomes $Pr(k_{A_0}, \ldots, k_T \mid \Omega^d_{A_0}, \text{type} = h)$. The likelihood contribution for the ith woman, $L_i(\Theta)$, is then equal to a weighted average of these terms, where the weights are the sample fractions $\mu_1, \ldots, \mu_H$, that is,

$$L_i(\Theta) = \sum_{h=1}^{H} \mu_h \prod_{t=A_0}^{T} Pr(k_{it} \mid \Omega_{it}, \text{type} = h), \tag{9}$$

where Θ is the vector of utility parameters (α, β, γ), wage and budget parameters (a, b) and error structure parameters (Σ), augmented to include the type probabilities μ.

Another issue is measurement error in earnings. The presence of substantial measurement error in the wage data is important, because the profile of the choice-earnings combinations k_t, and in turn our parameter estimates, will be strongly influenced by wage outliers. To attenuate this influence, we assume that wages are measured with error, that is,

$$y^r_t = \tilde{y}^r_t \exp(\epsilon^r_t + u^r_t), \qquad r = p, f,$$

where $u^r \sim \mathcal{N}(0, \eta^2_r)$, $E(u^p_t u^f_t) = E(u^p_{t-s} u^f_{t-k}) = 0$, for all $s \neq 0, k$, uncorrelated with ε's for all t; and $\tilde{y}^r_t$ is the deterministic component of the wage function as in equation (4). The parameters η_p and η_f will be estimated along with the other structural parameters included in Θ.

The identification of wage parameters relies neither on the logarithmic form of the earnings functions nor on the quadratic-in-experience specification: the wage parameters are identified from data on participation and earnings. From the alternative-specific value functions, all structural parameters can be estimated except that the utility parameters α cannot

[16] Standard errors are computed using a scoring method that calculates an outer product approximation to the hessian with numerical first derivatives. See Keane and Wolpin (1994) for a discussion of the properties of this estimation procedure.

be distinguished from the budget parameters b. Our identifying restriction is then $b_1 = b_2 = b_3 = 0$. This implies that α_1 and α_2 should be interpreted as gross costs of work (normalized to dollars), while α_3 and α_4 measure the value of children net of the goods cost of children.[17]

V. Data

The data used for estimation come from the NLS of Young Women, 1968–91. The original sample contains 5,159 women between the ages of 14 and 24 in 1968, when the survey was first conducted. These women had been interviewed 16 times over the 24 years between 1968 and 1991. We analyze two samples. Because many economic theories of household production, female labor supply, and fertility (Gronau 1973; Heckman and MaCurdy 1980) are meant to describe the behavior of women in long-term relationships with the same partner, the first sample (labeled as "always-married") consists of white women aged 19 or more at marriage, who were married only once and whose spouse was present at each interview. Conditioning on long-term marriage, however, raises the possibility that selection bias affects the estimates in this sample. That is, if the unobservables affecting marriage stability are also correlated with the fertility and participation decisions, then the always-married sample, which is selected on the basis of long-lasting relationships, will lead to biased estimates. For this reason, we analyze a second sample ("ever-married"), which includes all white women aged 19 or more who had a partner present at any interview regardless of their past marital status.[18] For each woman in these two samples, we use data on (i) her sociodemographic characteristics (age, age at marriage, schooling); (ii) her entire birth history; (iii) whether or not she worked in each year (but two) of the sample period, and, if she worked, her annual hours of work; (iv) the market wage rate in 1990 dollars for each year in which she worked; (v) some pre-1968 labor market information (but not wages); and (vi) her husband's earnings. We shall see that the basic observable characteristics, wage profiles, participation and fertility patterns, and choice distributions

[17] As long as the budget parameters are linear in p_t, f_t, and N_t (as in eq. [2]), they are not distinguishable from their respective psychic values in the utility function, α_1, α_2, and α_3 and α_4.

[18] Notice that the ever-married sample comprises all the women in the always-married sample plus the women who had more than one partnership or whose marriage ended during the survey period. To analyze labor force and fertility dynamics, we further restrict this second sample to women with at least 2 consecutive years of data on labor force participation and children. This restriction is not needed for the always-married sample.

are remarkably similar for women in the two samples. Our structural estimation is therefore performed on the always-married sample only.[19]

Before describing the data in detail, we ought to stress the importance of the sample selection problem and consider what biases might be introduced into the analysis by focusing on the subset of women in the always-married sample. The estimates derived from this sample are likely to be of general interest and thus contribute to the literature of female labor supply and fertility. This is because they are structurally obtained from a model grafted on long-standing economic theory (which requires undisrupted marriages) and because they are the first ones to respect the joint dynamic nature of fertility and labor supply decisions. But the similarity across the always-married and ever-married samples cannot be taken as evidence that the marital status decision is uncorrelated with either labor force participation or fertility (Johnson and Skinner 1988; Haurin 1989; Van der Klaauw 1996). The sample selection bias may come about in different ways. For example, if the unobservables that produce long-term relationships make women more desirable in the labor market (e.g., good communication and conflict management skills), then the always-married sample will be disproportionately represented by high-productivity women. In this case, we may expect the estimated labor-market-decision parameters to overstate their true values. If instead other women's unobservables lead to successful (long-term) marriages, such as preference for nonmarket activities, family commitment, and high specialization in household production, our sample will have a disproportionately large group of women with lower earning potentials and possibly yield downward biased estimates of the parameters related to the labor market decisions. Moreover, had these same unobservables been correlated with a stronger taste for children, then high-fertility women will be overrepresented in the always-married sample. While an overrepresentation of low-fertility women will occur if long-term unions are more likely to be characterized by women with a marked preference for child "quality" rather than quantity. In any case, marriage, fertility and participation are inextricably linked, and our always-married sample is bound to suffer from some sort of selection bias. To improve our understanding of such a bias and better assess our estimates, we shall introduce different sources of unobserved heterogeneity, which are meant to sort women out in their attitude toward children and work (see Sec. VI).

[19] Our estimation results, however, cannot be easily generalized. For instance, we have also studied a sample of always-married black women and found that their behavior was very similar to that of the corresponding group of white women (Francesconi 1995). But a comparison of the always-married black women sample with a sample of ever-married black women shows distinct behavioral differences, especially in fertility patterns. Because of the greater potential for sample selection bias, we limit our analysis to white women's samples.

A complete employment history is not readily available for all women even if they have no missing data. The first reason is because in the 1975 and 1977 interviews, which both followed a noninterview year, the current (last) job questions refer back to the date of the last interview, while the summary weeks and hours questions are asked about the last 12 months. To fill in the gaps of the work history, we use information on weeks and hours worked in the 12 months before interview, and employment data between missing interviews. With this information it is possible to construct a reasonably accurate annual employment profile for each woman (see Eckstein and Wolpin 1989*a*).[20]Another reason is because the NLS collected only limited information on work experience prior to the first survey. In the 1968 interview, women who were not currently working were asked the exact year that they last worked. Responses were coded into broad employment categories, whereby the distinction between part-time and full-time labor force status prior to 1968 is not possible. For all years worked before 1968, therefore, a woman is assigned to the labor force state in which she was actually observed in her first year of work.[21] Women who were instead working in 1968 (88 women in our estimating sample) were only asked about their date of entry into the labor market. Thus total experience is known but sector-specific experience is not. We then used the same assignment rule as before: a woman is assigned to either of the two employment sectors in which she was observed in 1968 for all periods worked before that year, under the assumption of no transition across labor force state and no career interruption.

There are 1,783 women in the ever-married sample, and 765 women in the always-married sample (43% of the total). Not surprisingly, their distributions by number of years in the data differ by sample (see table

[20] Assignments to the different employment states for 1973 and 1975 were made sequentially as follows. First, a woman who did not change labor force status over the entire sample period was assigned to the same labor force status for those years. Second, a woman who did not change labor force status in the 3 years adjacent to 1973 and 1975 was assigned to the same labor force status for those years. Third, a woman who did change labor force status in the 3 years adjacent to 1973 and 1975 was assigned to the labor force status of the subsequent nonmissing year for those years (assignments made with information on the preceding nonmissing year produced almost identical results). Fourth, all cases with imputed labor force states were matched with the information on hours worked in the reference week of the 1973 and 1975 interviews. Women with conflicting records were dropped. Another alternative is to account for missing data in estimation. The contribution to the likelihood function for women with missing labor force status in 1973 and 1975 needs to be modified by summing the probabilities of each possible time path of part- and full-time work experience. The estimation results are similar to those reported in this study, but the optimization routine is more time-intensive.

[21] For simplicity, we assume that there is no transition between part- and full-time work in the pre-1968 years.

Table 1
Distribution of Women by Number of Periods in the Two Samples

Number of Periods	Always-Married		Ever-Married	
	Number of Women	Cumulative Proportion	Number of Women	Cumulative Proportion
2			3	.002
3			16	.011
4			102	.068
5			128	.140
6			180	.241
7			133	.315
8			130	.388
9	5	.008	101	.445
10	3	.012	88	.494
11	14	.030	86	.542
12	0	.030	74	.584
13	29	.068	29	.600
14	20	.094	20	.611
15	1	.095	1	.612
16	68	.184	68	.650
17	9	.196	9	.655
18	124	.358	124	.725
19	58	.434	58	.757
20	78	.536	78	.801
21	73	.631	73	.842
22	65	.716	65	.878
23	53	.786	53	.908
24	164	1.000	164	1.000

1). In the ever-married sample, where women are included as long as their partner (whether first husband or not) is present, almost 50% of the women have 10 years of data or less, and only a quarter of them have 18 years of data or more. In the always-married sample, about four-fifths of the women have at least 18 years of data, and more than one-fifth have all 24 years of data. Only 1% of them have 10 years of relevant information or less. By construction, because of the longer length of their marital unions, women with 13 or more years of data can contribute only to the always-married sample.

Labor market involvement of wives is distinguished into three states: nonmarket work, part-time work, and full-time work. The definition of labor force status is based on annual hours of work (Corcoran et al. 1983; Nakamura and Nakamura 1983): working less than 500 hours per year defines nonwork, working between 500 and 1,500 hours defines part-time work, and working more than 1,500 hours defines full-time work.[22] Table 2 shows the distribution of labor force status by female age and by number

[22] Studies that define labor force status with information on usual weekly hours of work include Jones and Long (1981), Fraker and Moffitt (1988), Blank (1989, 1994), and Lehrer (1997). Ermisch and Wright (1993) use a woman's own definition of whether she works part-time or full-time.

Table 2
Labor Force Status by Age Group and Number of Children

	Always-Married				Ever-Married			
	OLF	PT	FT	*N*	OLF	PT	FT	*N*
All women	.438	.182	.380	15,162	.433	.198	.369	22,701
Age groups:								
19–24	.353	.190	.458	1,845	.368	.201	.431	3,516
25–29	.479	.173	.348	3,540	.491	.192	.317	5,458
30–34	.505	.174	.321	3,805	.501	.190	.309	5,712
35–39	.425	.191	.384	3,644	.403	.206	.391	5,126
40+	.350	.192	.458	2,328	.325	.207	.468	2,889
Number of children:								
0	.310	.173	.517	5,489	.324	.181	.494	8,313
1	.485	.194	.321	5,935	.488	.209	.303	8,994
2+	.549	.178	.274	3,738	.509	.206	.285	5,394

Note.—OLF = out of the labor force (nonwork); PT = part-time work; FT = full-time work; *N* = number of person-year observations. Rows may not add up due to rounding.

of children in the two samples. Of the 15,162 annual observations in the always-married sample, 18% are part-time employment periods, and 38% are full-time employment periods. The figures are remarkably close to those found with the ever-married sample, where the proportion of part-time employment periods is just two percentage points higher. Similar statistics are in Nakamura and Nakamura (1983) and Fraker and Moffitt (1988). Both samples show that the age-employment profile is nearly constant in part-time work but U-shaped in full-time work: women working full-time move to nonwork when aged 25–34 and then return to full-time employment. Approximately 50% of the women in both samples work full-time when they have no child. Most of the women, however, withdraw from full-time employment as the number of children increases and move into nonwork.

An important feature of the female labor market participation is persistence (Blank 1989, 1994; Eckstein and Wolpin 1989*a*). Our data confirm such a feature. The 1-period transition matrices in table 3 show that 79% of the women in the always-married sample (84% in the ever-married sample) stay in the same labor force state between any 2-year period, 7% (6%) leave employment, 7% (5%) enter the labor market, and another 6% (5%) move between part-time and full-time work. Labor market persistence is high especially in nonwork and full-time work. At any given period, women in either of these states have an 85% probability (88% probability in the ever-married sample) of remaining in the same state in the following period. In comparison, part-time employment is a more transitional state, particularly in the always-married sample. A part-time worker has only 58% probability of working part-time in the subsequent period and is also more likely to change labor force state, with 17%

Table 3
Labor Force Transition Matrices

Labor Force State ($t-1$)	Labor Force State (t) OLF	PT	FT
Always-married:			
OLF:			
Row %	84.53	9.85	5.62
Cell %	37.43	4.36	2.49
PT:			
Row %	25.28	57.72	17.00
Cell %	4.63	10.56	3.11
FT:			
Row %	7.05	8.97	83.98
Cell %	2.64	3.35	31.42
Ever-married:			
OLF:			
Row %	87.57	7.97	4.46
Cell %	38.48	3.50	1.96
PT:			
Row %	18.88	68.98	12.14
Cell %	3.67	13.40	2.36
FT:			
Row %	5.69	6.75	87.56
Cell %	2.08	2.47	32.08

NOTE.—OLF = out of the labor force (nonwork); PT = part-time work; FT = full-time work. There are 14,397 and 20,918 transitions in the always-married sample and ever-married sample, respectively.

probability of moving into a full-time job and 25% probability of moving into nonwork.

Table 4 reports descriptive statistics for the variables used in estimation. To forecast the husbands' earnings, equation (3) is specified by a logarithmic regression of husband's earnings on a constant; a linear and a quadratic term for the wife's age, age at marriage, and schooling; and a set of schooling-age at marriage interaction dummies (women with less than 12 years of schooling and aged 21 or less at marriage are the reference category). The linear estimates of such regressions are reported in table A1, which show that both samples generate very similar results. We assume that each woman forms expectations about her husband's future earnings in a manner that is consistent with this equation. In the always-married sample, part-time earnings are available for 1,615 person-periods (or 58% of all part-time employment periods) and full-time earnings are available for 3,268 person-periods (57% of all full-time employment periods) in the case of white women. In the other sample, the figures are 2,617 person-periods (58%) and 4,689 person-periods (56%) for part- and full-time work, respectively. Part-time earnings are more dispersed than full-time earnings, regardless of the sample. Table A2 presents the results from ordinary least squares regressions of log real annual earnings on part- and full-time experience and schooling using specification (4). Al-

Table 4
Summary Statistics of the Variables Used in Analysis

	Always-Married		Ever-Married	
	Mean	SD	Mean	SD
Age	32.464	6.332	31.677	6.414
Age at marriage	22.742	2.285	22.266	2.643
Schooling	13.374	1.948	13.107	1.895
Years of part-time experience	2.515	2.666	2.604	2.640
Years of full-time experience	4.274	4.511	4.223	4.229
Total number of children*	.932	.874	.913	.859
Total number of children†	1.460	.652	1.441	.635
Total number of children aged:†				
Less than 3	1.184	.412	1.192	.420
3–6	1.180	.408	1.186	.414
6 or more	1.538	.674	1.529	.668
Husband's expected annual earnings‡	23,889	6,578	23,791	7,708
Wife's part-time annual earnings§	7,794	6,488	7,627	6,223
Wife's full-time annual earnings§	16,401	8,095	16,073	8,123
Person-year observations	15,162		2,154	

* Computed on all person-year observations.
† Computed on person-year observations with positive number of children only. For women in the always-married sample the number of person-year observations are as follows: all children—9,673; children aged less than 3—2,141; children aged 3–6—2,219; and children aged 6 or more—6,448. The corresponding numbers for the ever-married sample are 14,388; 3,280; 3,368; and 9,401.
‡ In 1990 dollars.
§ In 1990 dollars. Computed on person-year observations with positive private-sector-specific earnings. For women in the always-married sample the number of person-year observations are 1,615 and 3,268 for part-time and full-time work, respectively. In the ever-married sample, the corresponding number of person-year observations are 2,617 and 4,689.

though these estimates are not selectivity corrected and assume a zero covariance between ϵ^p and ϵ^f, they show some relevant features of the data. First, schooling effects are strongly positive in both employment sectors. Second, own-experience effects are concave in full-time employment and convex in part-time employment. Third, cross-experience effects are weaker than (or go in the opposite direction of) own-experience effects. A 1-year increase in full-time experience has a small effect on part-time earnings, while a 1-year increase in part-time experience reduces full-time earnings for all women with average part-time experience.[23] In general, the larger size of the ever-married sample leads to smaller estimated standard errors. But the point estimates from the two samples are

[23] Similar results are obtained when we control for wife's age, age at marriage, number of children by age, and husband's earnings.

remarkably close to each other. The hypothesis of equality of the two sets of coefficients cannot be rejected at standard levels of significance.

As a way of describing the participation patterns, we perform a three-state multinomial logit regression, which includes the variables in table 4 as regressors and distinguishes three groups of children by age (less than 3, 3–6, and 6–18). The estimates, reported in table A3, are consistent with those found in other studies (see Lehrer 1997). Higher levels of schooling and work experience are associated with increased likelihood of current labor force participation, both part- and full-time. However, experience in part-time jobs is associated with a considerably larger level of part-time employment than full-time employment, with full-time experience having an opposite relationship. Women whose husbands have higher earnings are significantly less likely to participate in either employment sector. Women with young children (aged 0–6) have reduced participation in both employment sectors, while a higher number of children aged 6–18 reduces participation in part-time employment only but increases full-time participation. All these findings clearly emerge in both samples. An interesting difference across samples is that age at marriage does have an effect in the always-married sample (a higher age at marriage reduces the likelihood of full-time employment and increases the likelihood of part-time employment) but does not in the ever-married sample. This may partly reflect the lower age at marriage for women in the latter sample.[24]

Estimates of fertility and labor market behavior of always-married women do not necessarily generalize. The timing and number of births, the attachment to the labor market, and the complexity of intrafamily resource allocation are likely to differ between always-married women and other women. Nonetheless, our descriptive analysis reveals that women who are continuously married throughout the survey period and other (married) women show remarkably similar patterns in their fertility and labor market behavior. We believe, therefore, that the results derived from the always-married sample and presented in the following sections are possibly of general interest.[25]

[24] The two samples have been analyzed along other dimensions and are generally found to produce comparable results. For example, the hypotheses that birth hazard rates and choice distributions obtained from the two samples come from the same population of women cannot be rejected at conventional levels of statistical significance. We discuss both choice distributions and birth hazard rates in the next section.

[25] To assess the extent of sample selection further, we have compared the sample of always-married women to the sample of 1,018 "not-always-married" women (i.e., ever-married minus always-married, 1,783 − 765 women) over the period in which they have lived with a partner. Again, we find similar results for these two groups of women in terms of their labor force status by age group and number of children, labor force transitions, basic observable characteristics, wage equation estimates, and participation and fertility patterns. Although this similarity may

VI. Results

For the purpose of estimation, the population of always-married women has been divided into three types ($H = 3$), who differ in their attitudes toward children and work (see the discussion on sample selection in Sec. V).[26] Specifically, unobserved heterogeneity arises in the utility function (through α_3 and α_4) and in the earnings functions (through a_0^p and a_0^f), and μ_1, μ_2, μ_3 are the three type proportions.[27] We use three standards to assess our results: the plausibility of the parameter values (as compared to other studies in this literature), the ability of the model in replicating the observed choice distribution (other tests have also been performed), and out-of-sample fit.

A. Parameter Estimates

Table 5 provides estimates of the structural parameters and associated standard errors. Our parameter estimates seem to be comparable in magnitude to others available in the literature. For example, an additional year of schooling increases women's full-time earnings by 8.4% and part-time earnings by 7.6%. Eckstein and Wolpin (1989*a*) estimate a return to schooling of 5% for an older sample of women observed between 1967 and 1982. Van der Klaauw's (1996) estimate, for a sample of women aged 12–19 in 1968 and observed between 1968 and 1984, is of 9.8%. Using a sample from the NLS of Young Women observed between 1968 and 1973, Moffitt (1984) reports an estimate of 8.7%.

Some of the utility parameter estimates are also consistent with those

increase our confidence in the estimates presented below, it still does not eliminate the possibility of sample selection. In fact, we do find some differences in the timing of first (and higher order) births and in the pattern of the choice distribution between the always-married sample and the not-always-married sample. Moreover, none of our comparisons deal with the increasingly large group of never-married mothers (McLanahan and Sandefur 1994).

[26] We have also experimented with $H = 2$, 4, and 5. Increasing the number of types from two to three reduces the likelihood value by 37 points. Given that five extra parameters are estimated, this improvement in the likelihood is statistically significant at the 5% level ($-2\log L = 74.2$ while $\chi^2_{(5)}(0.05) = 11.07$). The introduction of four types ($H = 4$) leads, instead, to an insignificant improvement in the likelihood, i.e., ($-2\log L = 7.6$). A further increase in the number of types increases the likelihood value and deteriorates the model fit. The parameter estimates are, however, similar for the different values of H, when $H \geq 3$.

[27] The choice of parameters that are let to vary by type is arbitrary. We have performed several other estimations in which unobserved heterogeneity arises in different sets of utility and wage parameters. Most of the results and all the simulations remained unchanged (Francesconi 1995). The household discount factor is constant at 0.952, which corresponds to an annual rate of time preference of 0.05. A single iteration in the maximization of the likelihood function takes approximately 3 minutes CPU time on an IBM RISC-6000 machine.

Table 5
Maximum Likelihood Estimates: Always-Married Sample

Parameter	Estimate	SE	Parameter	Estimate	SE
Utility function:			Earnings functions:		
α_1	−768.345	123.160	a_0^p (h = 1)	6.683	.116
α_2	−1,474.540	238.403	a_0^p (h = 2)	7.004	.144
α_3 (h = 1)	5,288.083	292.114	a_0^p (h = 3)	7.549	.133
α_3 (h = 2)	4,501.766	210.165	a_1^p	.027	.009
α_3 (h = 3)	4,084.244	116.330	a_2^p	−.0007	.0002
α_4 (h = 1)	−649.148	128.422	a_3^p	.032	.006
α_4 (h = 2)	−924.523	49.108	a_4^p	−.0013	.0003
α_4 (h = 3)	−1,047.714	90.132	a_5^p	.076	.029
β_1	−.018	.007	a_0^f (h = 1)	8.714	.089
β_2	−.061	.032	a_0^f (h = 2)	8.195	.105
β_3	−.029	.013	a_0^f (h = 3)	9.074	.094
β_4	−179.040	96.436	a_1^f	.010	.017
β_5	−542.811	204.723	a_2^f	−.003	.0008
γ_1	−112.431	49.681	a_3^f	.076	.008
γ_2	−38.785	7.365	a_4^f	−.00097	.00006
γ_3	−25.059	11.042	a_5^f	.084	.014
γ_4	−171.316	45.729			
γ_5	48.278	17.360			
γ_6	−12.627	3.412	Error structure:		
γ_7	−15.643	6.389	σ_p	.298	.061
γ_8	−89.149	26.546	σ_{pf}	.0083	.000012
γ_9	−358.703	149.274	σ_{pN}	−.0025	.0792
γ_{10}	36.081	9.754	σ_f	.184	.021
			σ_{fN}	−.0227	.0863
Type proportions:			σ_N	224.851	16.239
μ_1	.253	.039	η_p	.391	.067
μ_2	.260	.045	η_f	.239	.038
μ_3	.477	.060	log$L$	−11,257.603	

found by Eckstein and Wolpin (1989*a*).[28] First, participation decreases utility (both α_1 and α_2 are negative), particularly in full-time employment. Second, the value of goods consumption is reduced when women participate (β_1, $\beta_2 < 0$). This supports the notion that consumption and leisure—from either part- or full-time work—are complements. Third, the disutility of work (and particularly full-time work) increases with schooling ($\gamma_3, \gamma_8 < 0$), that is, schooling enhances home production. Fourth, there is little evidence of habit formation in the labor market, with the disutility of work increasing with the amount of prior experience ($\gamma_1, \gamma_2, \gamma_6, \gamma_7 < 0$). Finally, the marginal utility of children is lower when the mother participates (γ_4 and $\gamma_9 < 0$), but at a decreasing rate (γ_5 and $\gamma_{10} > 0$). The value of the first child to nonworking women is \$123 and \$322 more than to part-time and full-time workers, respectively. A second child would reduce the corresponding "penalty" to \$75 and \$286.

[28] Notice, however, that not all the results are directly comparable. This is because we estimate two separate sets of parameters for part-time work and full-time work and have a larger number of parameters on children.

Although the other estimates of the utility parameters cannot be readily compared with existing studies, they do provide a simple check on their plausibility. As expected, our estimates show that marginal utilities are positive and diminishing in children ($\alpha_3 > 0$, $\alpha_4 < 0$),[29] and vary across types substantially. For example, the (monetary) utility value of the first child (using only α_3 and α_4) is estimated to be \$4,600, \$3,600, and \$3,000 for type-1, type-2, and type-3 women, respectively. The number of children consistent with utility maximization (using α_3[type = 1, 2, 3], α_4[type = 1, 2, 3], β_3, β_4, β_5, γ_4, γ_5, γ_9, and γ_{10}) depends on the mother's employment sector and type. For type-1 women, it is 4.1 if they work part-time and 3.6 if they work full-time. The numbers decrease to 1.9 and 1.6 for type-3 women working part- and full-time, respectively. Notice that the value of goods consumption is reduced not only when women participate but also when a birth occurs ($\beta_3 < 0$). Similarly, the disutility of work increases when a birth occurs ($\beta_4, \beta_5 < 0$), although a two-standard deviation change would reverse the sign of these parameters.

With respect to the wage, error structure, and heterogeneity parameters, we draw attention to three sets of results. First, full-time experience increases full-time earnings by 7.5% in the first year; at 20 years of continuous full-time experience, the additional year adds 3.8%. Thus, the estimated full-time earnings profile exhibits a mildly concave shape over the work cycle, peaking at about 40 years of full-time experience.[30] The first year of part-time experience increases part-time earnings by 2.6%. After that, each additional year increases part-time earnings by $2.7 - .0014 \times K^p$ percent, with peak earnings reached at approximately 19 years of experience (age constant). Cross-experience effects are smaller than own-experience effects in the full-time sector, but are higher in the part-time sector. For example, at any level of part-time experience, 5 years of full-time experience increase part-time earnings by 13%; while 5 years of part-time experience increase the corresponding full-time earnings by only 0.7% (full-time experience constant).

Second, earnings profiles differ by type significantly. Table 5 shows the estimated proportion of women of each of the three types (μ). Almost 50% of the women are of type 3, with the other 50% equally divided between type 1 and type 2. Table 6 shows the relative rankings of the three types. Notice that type-3 women, the largest group in the population, have the highest earnings profiles (in both employment sectors) and the lowest marginal utility of children. Type 1's have the lowest

[29] Wolpin (1984) finds a similar result for a sample of 30–50-year-old Malaysian women.

[30] By comparison, experience effects peak after 60 years in Eckstein and Wolpin (1989*a*) and at 140 years in Van der Klaauw (1996). Hotz and Miller (1988), who do not include work experience in the wage equation, find that their estimated wage profile peaks much earlier, at about 44 years of age.

Table 6
Type-Specific Rankings

Rank Ordering	Type 1	Type 2	Type 3
Part-time earnings profile ($a_0^p(h)$)	3	2	1
Full-time earnings profile ($a_0^f(h)$)	2	3	1
Direct marginal utility of children ($\alpha_3(h)$ and $\alpha_4(h)$)	1	2	3

earnings profile in the part-time sector but the highest marginal utility of children, while type-2 women have the lowest full-time earnings profile and rank second elsewhere. Some patterns clearly emerge: type 3's display a comparative advantage in market work, type 1's have a strong preference for children (and may have an advantage in rearing them), type 2's do not display either a particular advantage or a particular disadvantage, except for a disadvantage in full-time employment. It therefore appears that our sample of always-married women is disproportionately composed of high-productivity, low-fertility women. Following our discussion on sample selection in Section V, it is hard to establish the direction of the bias introduced by our sample selection criteria, although this overrepresentation may possibly lead to an upward bias (in absolute value) of the (utility and wage functions) estimates related to the participation decisions and to a downward bias (in absolute value) of the estimates related to the fertility decision. We have, however, no additional information to assess the extent of such a bias.

Third, part-time wages are reported with the most error. Measurement error of part-time earnings is about 60% greater than that of full-time earnings (table 5). The proportion of the total wage variance due to measurement error is also large: measurement error accounts for 63% and 62% of the total (log) part-time and full-time wage variance, respectively. The correlation of earnings in the two employment sectors is positive and close to 15%. The estimated covariance between ϵ^n and either of the two wage residuals is estimated to be negative but always not significantly different from zero. Although well determined, the estimate of σ_n is small: doubling its value does not change the type rankings of the marginal utility of children shown in table 6.

In sum, most of our results appear to be reasonable in comparison with the existing literature on female labor supply and fertility. The earnings profiles in the full-time employment sector are, however, more concave than those found in other structural models of female labor force participation (Eckstein and Wolpin 1989*a*; Van der Klaauw 1996). This is perhaps the result of including fertility in the choice set. Furthermore, the distinction between part-time work and full-time work allows us to identify some of the sources of labor force state persistence. The observed

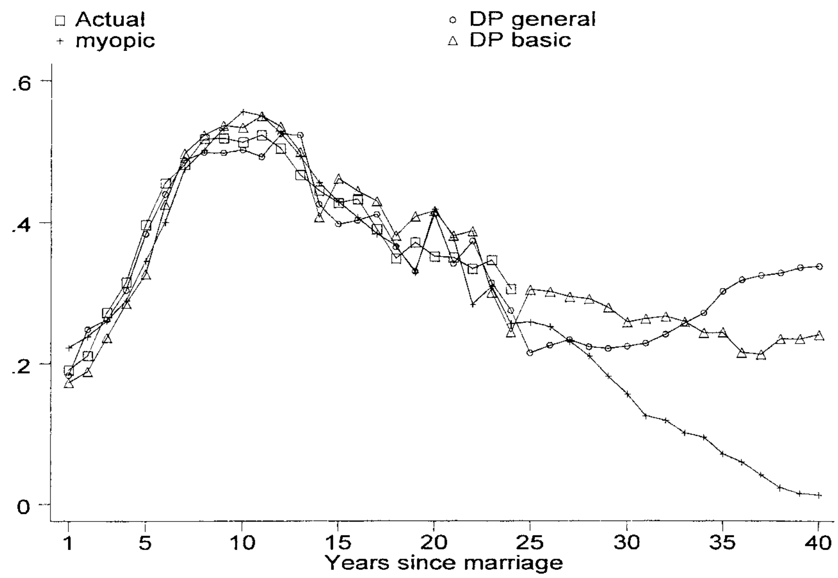

FIG. 1.—Nonwork and no child

persistence in full-time employment seems to be due to the large experience effect on full-time earnings, even though the disutility of work increases with full-time experience (see Sec. VII). The lower persistence in part-time employment is mainly due to the small experience effect on part-time earnings.

B. Model Fit

As shown in table A4, there are no statistically significant differences between actual and predicted transition matrices. Figures 1–6 show the fit of the model (with label "DP general") to the actual choice distribution by year of marriage. For comparison purposes we also show the predicted choice distribution of a static model ("myopic"), whose only difference with the DP-general model is that the discount factor is set to zero, and of a model that does not introduce unobserved heterogeneity, that is, it estimates the dynamic-programming model for one type of women only ("DP basic"). From the graphs, it is difficult to distinguish the fit of the three models, although it is clear that they all perform reasonably well in replicating the main features of the data.

Table 7 presents within-sample χ^2 goodness-of-fit test statistics.[31] The first column refers to the DP-general model whose estimates are presented

[31] The χ^2 statistics computed here have not been adjusted for the fact that the parameters of the model have been estimated.

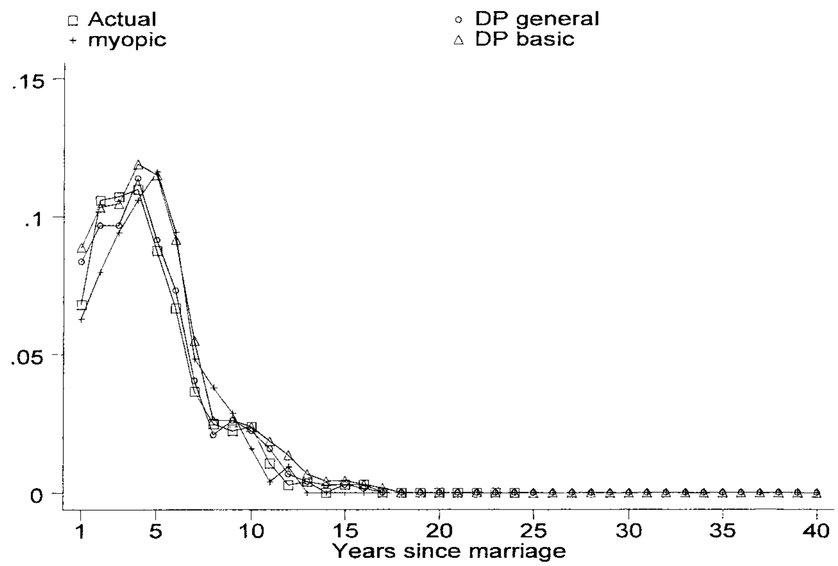

FIG. 2.—Nonwork and child

in table 5. The model performs extremely well in replicating period-by-period individual behavior. Only in three of the 24 years of data is the hypothesis of equality of the choice distributions rejected at the 5% significance level. Using this metric, the fit of the other two models is less satisfactory. Table 7 shows that the DP-myopic model fails to reproduce the choice distribution at the beginning of the period, while the DP-basic model cannot fit the individual choice data over most of the optimization period. This suggests that the variations by type play a considerable role in replicating the fertility and participation decisions of this sample of women. Other tests of model fit have been performed but are not presented because of space limitations (see Francesconi 1995). For example, we find that birth hazard rates—predicted with the estimates of the DP-general model—closely replicate the pattern of the actual rates.[32] This suggests that the model can predict the timing (and spacing) of birth events very accurately. Furthermore, the DP-general model can precisely replicate the observed age (since-marriage) profile of earnings for the two types of employment: mean predicted earnings lie always inside the 2-standard-deviation band.

[32] The birth hazard rates show an inverse U-shaped pattern, with the probability of having a child rising sharply from 0.02 at the beginning of the marriage to 0.22 in the sixth year and gradually declining thereafter.

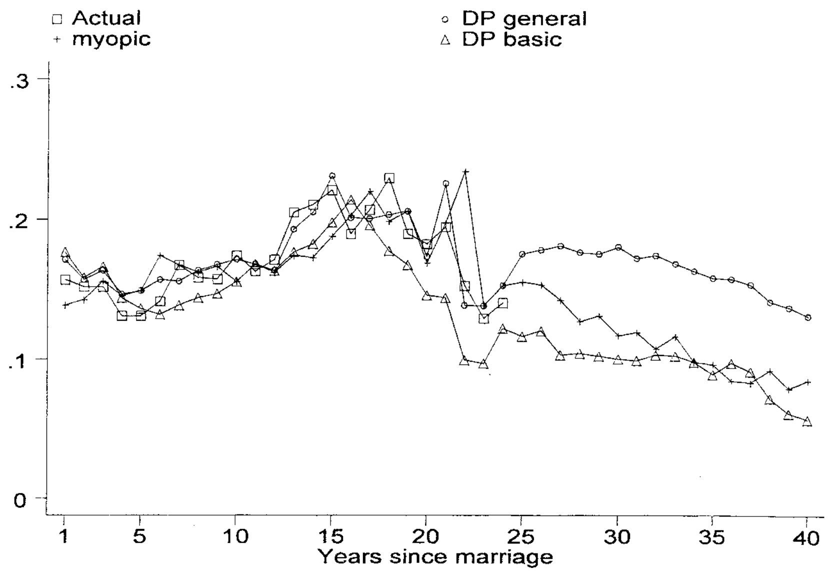

FIG. 3.—Part-time work and no child

C. Out-of-Sample Fit

Figures 1–6 also show the forecasts of the three models beyond the actual data, through 40 years of marriage.[33] All models have identical, reasonable fertility forecasts, with no birth predicted in the out-of-sample period. The forecasts, however, diverge considerably in predicting the labor force participation decision. The myopic model predicts a sharp change in the participation patterns, with almost no woman out of the labor force and 90% of the women choosing full-time work by their fortieth year of marriage (corresponding, on average, to age 63). This is an unrealistic projection. Neither the basic model nor the general model generate such extreme predictions: they both predict that at least a quarter of women would be out of the labor market by the end of their working careers. But they differ in predicting the distribution of women into part-time and full-time work. The basic model forecasts that, at the fortieth year of marriage, the full-time and part-time participation rates will be 70% and 6%, respectively. The corresponding rates predicted by the general model are 50% and 13%.

We obviously cannot contrast the out-of-sample fits of the basic and general models with the NLS cohort data used in estimation: the oldest members will not be 63 years old (when their marriage duration is on

[33] The out-of-sample fit is performed on a simulated sample of 1,000 individuals.

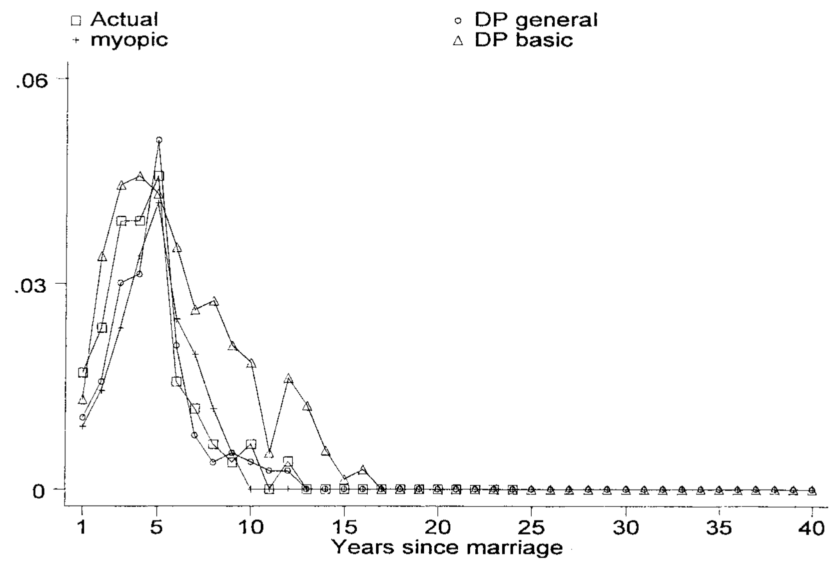

FIG. 4.—Part-time work and child

average 40 years) until 2007, and the youngest cohort members will be age 63 10 years later. However, some of the available evidence about the trends in the labor force participation of older women seems to be favorable to the general model rather than to the basic model. The full-time labor force participation predicted by the basic model is unreasonably high. For example, Peracchi and Welch (1994, table 1), using Current Population Survey data and the official definition of part-time work, show that the distribution of part-time and full-time employment of women aged 49–68 between 1968 and 1991 is 10% and 31%, respectively. The difference in full-time participation rate between the forecasts from the basic model and the CPS data cannot be due to cohort effects only.[34] Therefore, the DP-general model provides more convincing out-of-sample employment patterns than the other two models. By denying any forward-looking behavior (i.e., $\delta = 0$), the myopic model predicts an unreasonably large proportion of women in employment at the end of

[34] Peracchi and Welch (1994, table 2) show that the proportion of women aged 55 between 1987 and 1990 working full-time is 45%. As women become older, their full-time participation rate declines: the proportions of women aged 60 and 62 in the same period working full-time declines to 35% and 23%, respectively. In either case, the DP-basic model appears to overpredict full-time participation severely. Other evidence based on official statistics is not readily available. It is known, however, that the part-time participation rate tend to be larger among older women (Nardone 1986; Tilly 1991), casting doubts on the small rate predicted by the DP-basic model.

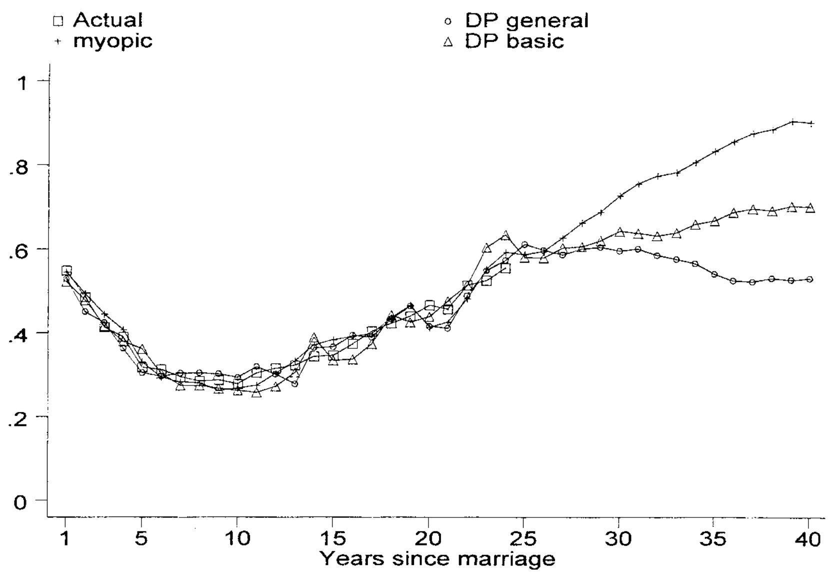

FIG. 5.—Full-time work and no child

their working life. However, the DP-basic model, which does not distinguish different types of women in terms of their preferences for children and work (i.e., $H = 1$), predicts an unreasonably large fraction of women in full-time jobs and a presumably too small proportion of women in part-time jobs by the end of their career.

VII. Discussion

A. The Shape of Participation and Fertility Profiles

Based on the estimated parameters of the DP-general model of table 5, we compute the probabilities of working by age and sector-specific experience, and the probabilities of having a child by age, sector-specific experience, and number of prior children.[35] For the sample of always-married women, we find that working probabilities increase with experience at each age. Conversely, for given experience in either sector the probability of work decreases with age (Eckstein and Wolpin 1989*a*). The employment profile over the fertile period can be traced out for different experience levels. For example, a woman with 5 years of full-time experience at age 25 who does not work full-time thereafter would decrease her probability of working full-time from 71% to 43% by age 40. If she works full-time for another 5 years out of the of 16 available (between

[35] For the sake of brevity, these figures are not reported. See Francesconi (1995).

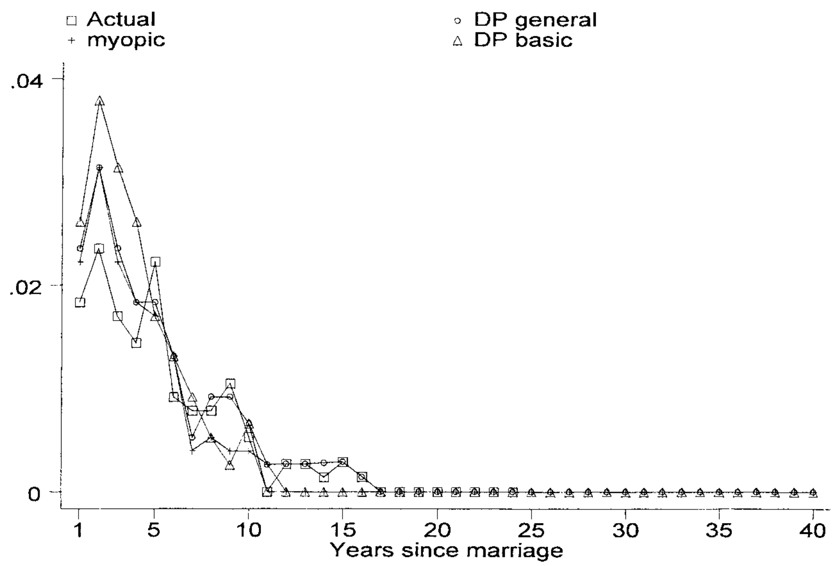

FIG. 6.—Full-time work and child

ages 25 and 40), her probability of working full-time will rise to 69% at age 40. If she works full-time in every subsequent year, her working probabilities will rise to 92%.

We also trace out the birth probability profiles over the fertile period (assumed to end at age 40) conditional on the number of children in the household at the beginning of marriage. We find that the probabilities of having a child decline with the number of children at any given age. Except when there are 2 children at the start of the marriage (which comprises only 12 of the 765 women in the sample), the pattern is for the birth probabilities to increase from age 25 to age 30 and to decline thereafter.[36]

B. Employment Simulations

To assess the effect of changing women's environment on their employment behavior, we perform several simulations. All experiments are evaluated in terms of their effects on the expected number of additional years of sector-specific work from age 30 to age 40. Table 8 reports the results from simulations concerning changes in women's schooling, in-

[36] This inverse U-shaped pattern is consistent with the birth hazard rates discussed in Sec. VIIB. See also Moffitt (1984) and Hotz and Miller (1988).

Table 7
χ^2-Goodness-of-Fit Tests of the Within-Sample Choice Distribution

Years since Marriage	Model Definition DP-General	Myopic	DP-Basic	Sample Size (df)
1	8.08	11.26*	10.61	765 (5)
2	11.68*	15.51*	8.97	765 (5)
3	5.60	12.51*	10.57	765 (5)
4	5.47	4.81	8.59	765 (5)
5	3.67	14.83*	21.93*	765 (5)
6	4.61	20.94*	16.71*	765 (5)
7	3.69	8.06	17.02*	765 (5)
8	3.72	6.72	14.69*	765 (5)
9	2.54	12.20*	31.32*	764 (5)
10	2.44	7.82	9.00	759 (5)
11	7.38	13.85*	14.04*	756 (5)
12	3.72	5.00	19.85*	742 (5)
13	10.39*	5.45	15.36*	742 (4)
14	4.12	7.63*	16.51*	713 (4)
15	2.68	6.61*	5.28	693 (4)
16	3.76	2.26	7.06	692 (4)
17	1.12	.60	5.28	624 (2)
18	2.63	3.74	11.63*	615 (2)
19	3.73	4.11	3.33	491 (2)
20	6.70*	8.09*	9.18*	433 (2)
21	3.34	1.56	7.54*	355 (2)
22	1.92	11.06*	10.10*	282 (2)
23	1.06	1.39	6.08*	217 (2)
24	.81	2.05	4.58	164 (2)

NOTE.—DP-general is dynamic programming model with unobserved heterogeneity (whose MLE estimates are presented in table 5); myopic is as DP-general but sets $\delta = 0$; DP-basic is as DP-general without unobserved heterogeneity; df = degrees of freedom of the test for DP-general model. χ^2 tests the difference between actual and predicted rates. $\chi^2 = \sum (N_P - N_A)^2/N_P$, where N_P = number predicted and N_A = number actual. The test is performed on cells with positive N_P only. $\chi^2_{(5)}(0.05) = 11.07$; $\chi^2_{(4)}(0.05) = 9.49$; $\chi^2_{(3)}(0.05) = 7.82$; and $\chi^2_{(2)}(0.05) = 5.99$. For the myopic model, df = 5 in rows 1–9, df = 4 in rows 10–11; df = 3 in row 12, and df = 2 in rows 13–24; DP-basic: df = 5 (rows 1–10), df = 4 (rows 11–16), df = 3 (row 17), df = 2 (rows 18–24).
* Denotes that actual and predicted rates are statistically different at the .05 level.

tercept,[37] and slope of full-time earnings. From the estimation, we found that schooling has two offsetting effects: in both employment sectors, it increases earnings (via a_5^p and a_5^f) and the disutility of work (via γ_3 and γ_8). Table 8A shows that increased schooling unambiguously increases the expected years of work, particularly in the full-time sector. This effect tapers off with experience: if a woman has never worked full-time, the elasticity of expected number of additional years worked with respect to schooling is about 2.8, but drops to 0.8 at 4 years of prior full-time experience.

For given work experience at marriage, the number of years worked

[37] The "mean" value reported in table 8B is computed as a weighted average of the estimated full-time earnings function intercepts, $a_0^f(h)$, reported in table 5, with weights given by the type proportions, μ_h, $h = 1, 2, 3$. All changes in table 8B are in levels rather than in logarithms.

Table 8
Expected Number of Years Worked Part-Time and Full-Time from Age 30 to 40 for Given Sector-Specific Experience
A. By Schooling Level

Sector-Specific Experience	Years of Schooling: Part-Time Work			Years of Schooling: Full-Time Work		
	8	12	16	8	12	16
0	1.719	3.643	4.511	1.926	4.721	7.318
2	3.604	4.236	5.054	3.849	5.465	8.024
4	4.893	5.145	5.920	5.254	7.221	9.128

B. By Earnings Intercept in Full-Time Sector ($a_0^f(h)$)

Sector-Specific Experience	Full-Time Wage Intercept ($a_0^f(h)$): Part-Time Work				Full-Time Wage Intercept ($a_0^f(h)$): Full-Time Work			
	−25%	−10%	Mean*	+10%	−25%	−10%	Mean*	+10%
0	3.545	2.815	2.133	.089	.064	2.709	5.108	11.0
2	4.986	3.057	2.562	.675	.112	3.921	6.833	11.0
4	5.902	3.420	3.742	1.021	.236	4.190	8.739	11.0

C. By Wage-Experience Slope in Full-Time Sector (a_3^f)

Sector-Specific Experience	Full-Time Wage Slope (a_3^f): Part-Time Work				Full-Time Wage Slope (a_3^f): Full-Time Work			
	0	.038	.076	.142	0	.038	.076	.142
0	2.919	2.431	2.220	.281	.845	1.720	4.976	11.0
2	4.343	2.718	2.461	.922	.812	2.013	6.841	11.0
4	5.137	3.953	3.738	1.381	.789	2.127	8.750	11.0

* Mean is a weighted average of the estimated full-time earnings function intercepts (see table 5), with weights given by the type proportions μ_h. All changes are expressed in wage levels rather than in logarithms.

by a married woman between ages 30 and 40 changes dramatically as the position of the full-time earnings profile changes (table 8B). Increasing the average profile by 10% implies that a woman, with any level of initial experience, will work full-time for all years from age 30 to age 40. Conversely, a downward shift of the average full-time earnings profile by 25% will drive women out of the full-time sector into either part-time work or nonwork.

Changing the wage-experience profile yields powerful effects on employment patterns (table 8C). Because full-time experience increases the disutility of work, a flat profile of wage and experience ($a_3^f = 0$) would lead to a negative experience profile. Halving a_3^f reduces future full-time employment by almost 80% for a woman with 4 years of full-time experience at marriage. But if a_3^f doubles, all women will work full-time in every year subsequent to age 30, independent of work experience at age 30. The elasticity of expected number of years worked part-time with respect to a_3^f is clearly smaller. In comparison, an increase in a_3^p (not shown here) will induce an increase in part-time participation but to a lesser extent than in the case of an increase in a_3^f. Full-time work is then only mildly substitutable with part-time work, while leisure seems to be more easily intertemporally substitutable with full-time work than with part-time work.

In sum, both education and full-time experience behave in the same way. They increase the disutility of full-time work but increase full-time earnings sufficiently to yield a positive schooling-participation profile and positive state dependence. However, there is little evidence that part-time work patterns mimic full-time work patterns. Attracting a larger number of women into the part-time employment sector would require substantial changes in women's socioeconomic environment. For example, the expected number of years worked part-time between age 30 and age 40 by a college graduate is only 6 months higher than the number of years worked full-time by a high-school dropout.

C. Fertility Simulations

We perform similar simulations to analyze the effect of changing women's socioeconomic environment on their fertility behavior. The simulations are evaluated in terms of their effects on the expected number of additional children from ages 25–40, given the number of children at marriage. The results are reported in table 9.

Increased schooling decreases the expected number of children substantially (table 9A). A downward shift of the full-time earnings profile (table 9B) increases fertility:[38] for example, a 25% reduction will increase the number of births from 2.5 to 2.9 for a woman who was childless at marriage and accumulates less than 5 years of part-time experience by age 40; for a woman with less than 5 years of full-time experience and no children at marriage, the same shift in the full-time earnings profile will yield an increase of the number of births from 2.3 to 3.5. Changes in the full-time wage-experience profile exhibit strong effects on the expected number of births (table 9C). A flat wage-experience profile increases fertility: a woman, who is childless at marriage and acquires less than 5 years of full-time experience by age 40, will expect to have 3.7 children at that age if $a_3^f = 0$. For the same woman, halving the slope of the log wage-experience profile in full-time earnings will lead to 3.0 children, while doubling the coefficient will lead to 1.8 children by end of the fertile period.[39]

[38] The table shows that no woman is predicted to have more than 5 years of full-time experience when $a_0^f(h)$, for all h, is set to zero. A similar prediction occurs when $a_3^f = 0$ in table 9C.

[39] Changing the experience profile of part-time earnings (not shown) has a more limited effect on fertility. This effect may work through two mechanisms. One mechanism has to do with the earnings technology and the substitutability of work with leisure, which is higher for part-time work than full-time work (see

Table 9
Expected Number of Additional Children from Age 25 to 40 for Given Number of Children
A. By Schooling Level

Number of Children	Years of Schooling: Less than 5 Years			Years of Schooling: More than 5 Years		
	8	12	16	8	12	16
Part-time work:						
0	3.656	2.752	2.214	3.498	2.769	2.178
1	2.353	1.870	1.559	2.275	1.818	1.711
2	.909	.578	.475	1.637	1.216	.659
Full-time work:						
0	4.036	2.587	1.903	2.915	2.242	1.677
1	2.314	1.989	1.441	2.205	1.528	1.280
2	2.061	.858	.626	1.464	.749	.399

B. By Earnings Intercept in Full-Time Sector ($a_0^f(h)$)

Number of Children	Full-Time Wage Intercept ($a_0^f(h)$): Less than 5 Years				Full-Time Wage Intercept ($a_0^f(h)$): More than 5 Years			
	−25%	−10%	Mean*	+10%	−25%	−10%	Mean*	+10%
Part-time work:								
0	2.904	2.461	2.452	2.184	3.020	2.520	2.142	1.970
1	2.687	2.263	1.536	1.357	1.994	1.533	1.361	1.273
2	2.250	1.658	1.120	.882	2.010	1.210	.852	.784
Full-time work:								
0	3.502	3.124	2.272	2.011	. . .	2.478	1.974	1.755
1	2.773	2.431	1.479	1.724	. . .	1.968	1.270	1.141
2	1.654	1.323	.944	.826	. . .	1.243	.793	.663

C. By Wage-Experience Slope in Full-Time Sector (a_3^f)

Number of Children	Full-Time Wage Slope (a_3^f): Less than 5 Years				Full-Time Wage Slope (a_3^f): More than 5 Years			
	0	.038	.076	.142	0	.038	.076	.142
Part-time work:								
0	3.096	2.949	2.489	1.989	3.187	2.486	2.118	1.869
1	2.863	2.503	1.543	1.302	2.016	1.712	1.382	1.256
2	2.314	1.997	1.117	.579	1.263	1.093	.876	.762
Full-time work:								
0	3.128	3.062	2.260	1.752	. . .	2.438	1.982	1.668
1	2.893	2.214	1.439	.931	. . .	1.622	1.235	1.109
2	1.964	1.538	.960	.623	. . .	.978	.767	.588

* Mean is a weighted average of the estimated full-time earnings function intercepts (see table 5), with weights given by the type proportions μ_h. All changes are expressed in wage levels rather than in logarithms.

D. Work Interruption or Part-Time Work?

It has long been argued that interruptions in women's career associated with the birth of children reduce female earning power (Weiss and Gronau 1981; Mincer and Ofek 1982; Corcoran et al. 1983; Moffitt 1984; Nakamura and Nakamura 1992; Belzil and Hergel 1999). This is because a withdrawal from the labor market results in the loss of current earnings and lowers the investment in human capital. Is part-time employment a viable alternative to work interruptions? Many advocates of public policies that encourage highly paid part-time work during childbearing years underpin their argument with the idea that part-time work is a bridge between nonwork and full-time employment (Blossfeld and Hakim 1997).

Using the estimates of our DP-general model (table 5), we can address this question by tracing out the lifetime utility derived from different employment strategies and fertility decisions. We assume that a woman has her first child in the fourth year of her marriage.[40] Four different employment strategies are compared. The first involves no interruption of full-time work over the 24-year period under analysis. The second strategy requires a withdrawal from the market for a year at the child's birth, followed by reentry into full-time job until the end of the period. The third strategy differs from the second strategy in that it involves part-time employment rather than nonwork in period 4, when the child is born. The last strategy requires a woman to switch from her full-time job to a part-time job when she has the child and to stay in the part-time sector until the end of the period.

In the top panel of figure 7*a*, we plot the present discounted value of utility (PDVU in $1,000) from these four employment strategies. The bottom panel of the figure plots the PDVU ratios of strategies 2–4 to strategy 1 (our base case). In figure 7*b*, we consider the effect of a second child in period 7 (3 years after the first child was born). The first and the fourth employment strategies are not altered. But strategy 2 now requires a woman to exit from the labor market for another year when the second child is born, while strategy 3 requires a woman to enter the part-time sector for that year only. Again, the top panel of the figure graphs the PDVU of the four strategies, and the bottom panel graphs the PDVU ratios with strategy 1 as the base. Finally, figure 7*c* illustrates the effect

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Sec. VIIB). The other mechanism is purely based on preferences. The disutility of part-time work increases with the number of children by a smaller extent (i.e., $\beta_4 < \beta_5$ in absolute value); and increases in the level of part-time experience have a smaller impact on future part-time participation than increases in full-time experience have on future full-time participation (i.e., $\gamma_1 < \gamma_6$, and $\gamma_2 < \gamma_7$).

[40] Life table estimates show that 35% of the women in the always-married sample have a child by their fourth year of marriage. This figure is slightly higher in the ever-married sample. It reaches almost 50% for the not-always-married sample of women.

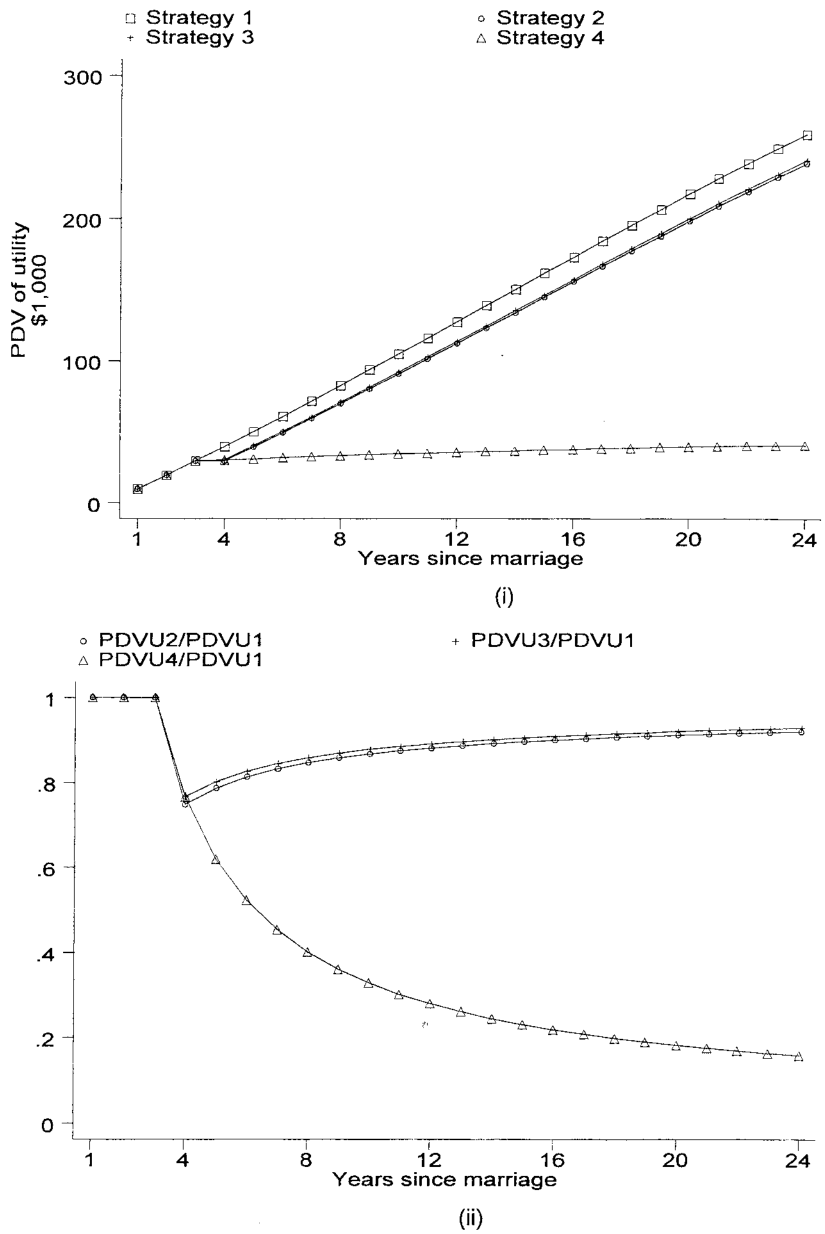
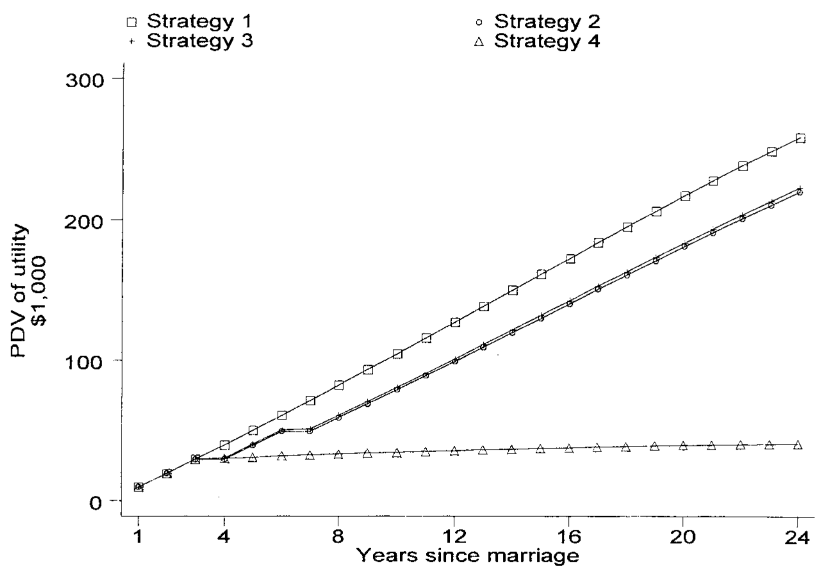


FIG. 7a.—Work interruption or part-time work? (One child)



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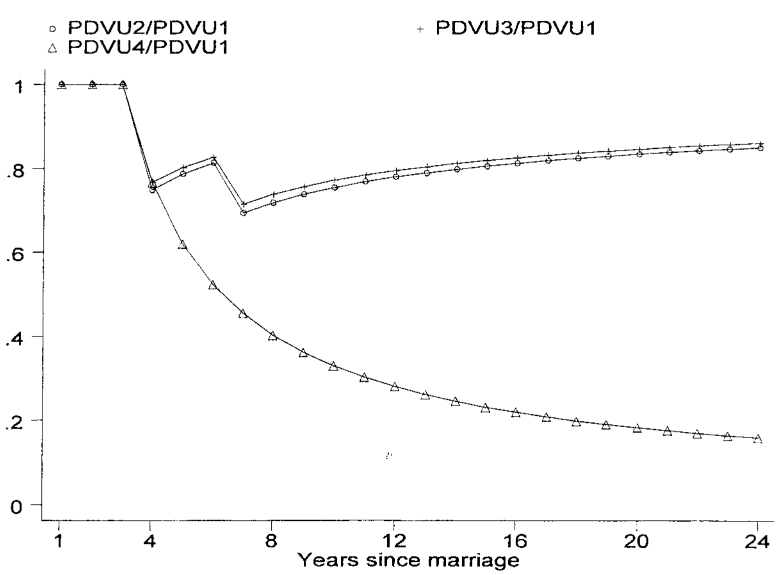
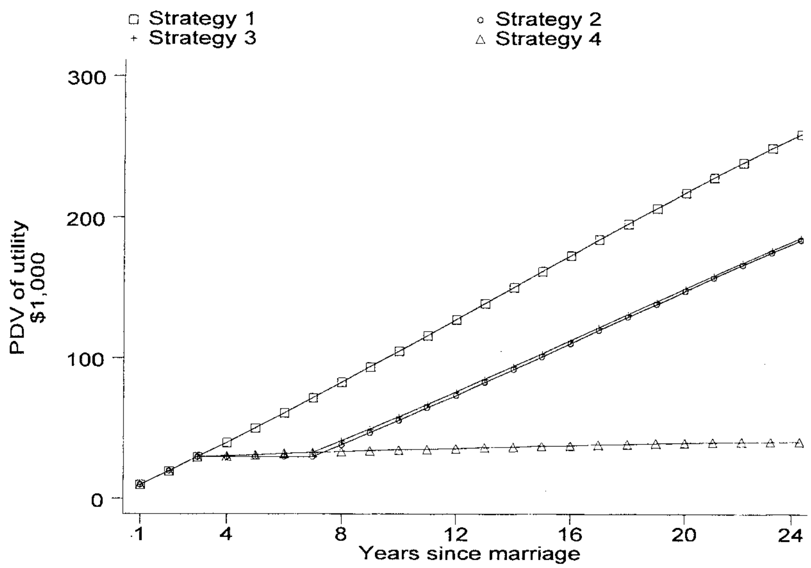


(iv)

FIG. 7*b*.—Work interruption or part-time work? (Two children)



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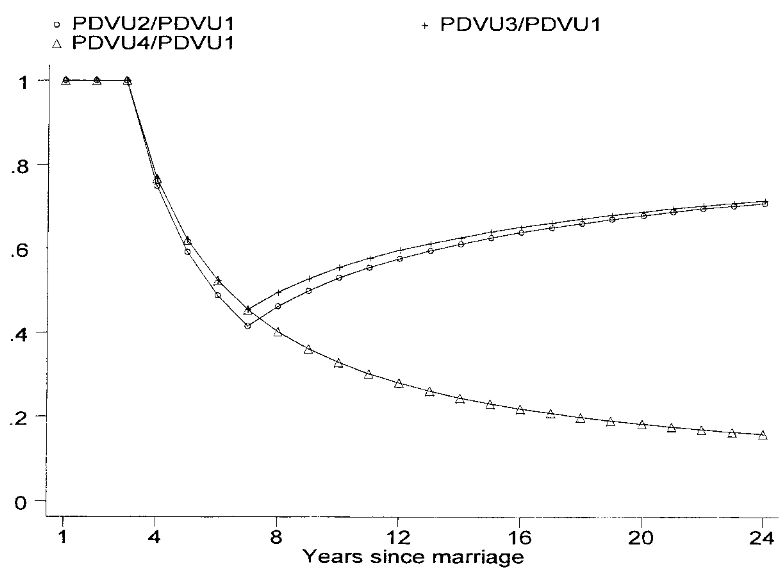

(vi)

FIG. 7*c*.—Work interruption or part-time work? (Longer withdrawal from full-time employment).

of a 5-year withdrawal from the full-time sector into either nonwork (strategy 2) or part-time work (strategy 3). The other two strategies are identical to the corresponding strategies of figure 7*a*.

Figures 7*a*–*c* clearly show that continuous full-time work yields the highest earnings profiles. Work interruptions due to the birth of a child reduce current utility substantially (fig. 7*a*): a 1-year interruption decreases utility by 25% (from $40,000 to $30,000) in the year of the childbirth as compared to a noninterrupted full-time work career. We find a strong "rebound" effect upon reentry into full-time employment: the present discounted value of forgone utility is $20,000, that is, less than 8% of the PDVU from strategy 1 over the entire period (or 0.4% per year). More frequent and longer work interruptions (figs. 7*b* and 7*c*) have more durable effects on the woman's lifetime utility. Although the rebound effect upon reentry into full-time work is still sizable, such interruptions shift the full-time earnings profile at a lower absolute level, and lifetime utility cannot attain the same level as that achieved under strategy 1 because the woman faces a finite horizon. Finally, the figures show that a reentry into full-time employment from part-time employment does not lead to a larger rebound effect over the life cycle. It is only when the woman opts for a long withdrawal (fig. 7*c*) that we observe a short-term utility loss as compared to part-time employment. This loss reaches $3,000 by year 7, but it vanishes by the end of the period. Hence, it appears that part-time employment does not provide a viable bridge between nonwork and full-time employment for women who face interruptions in full-time employment due to childbearing. Some caution toward policies that encourage part-time employment during the childbearing/child-rearing period is warranted.

VIII. Conclusions

In this article we have estimated a joint dynamic structural model of labor supply and fertility decisions of married women using data from the NLS of Young Women, 1968–91. The theoretical framework integrates previous work by Heckman and MaCurdy (1980), Weiss and Gronau (1981), Moffitt (1984), Hotz and Miller (1988), Eckstein and Wolpin (1989*a*), and Angrist and Evans (1998) by jointly modeling and estimating, for the first time, labor supply and fertility decisions with the specification of six alternative choices. We distinguish two employment sectors, part-time and full-time work: they are found to be different both in terms of their pecuniary (wage) and nonpecuniary (utility) returns, and in terms of their transferability of human capital. Self-selection among the six choices, based on expected future utility and unobserved heterogeneity in preferences for children and earning ability, is shown to be an important element in fitting the data on the fertility and participation choices of this specific cohort of women.

Our structural estimates are derived from a sample of continuously married women. But a comparison of such a sample with another sample of ever-married women (regardless of their marital history) reveals that women in the two samples are remarkably similar in their labor market and fertility behavior. Although the potential for sample selection bias cannot be ruled out, our estimates are nonetheless of general interest. The results show that while the disutility of work increases with own experience, which by itself would lead to negative state dependence, the observed persistence in full-time employment is due to the large positive own-experience effect on full-time earnings. The observed lower persistence in part-time employment is mainly due to the small own-experience effect on part-time earnings. The behavioral effect of schooling, which enters both the utility function and the earnings equations, is similar to that of full-time experience. By increasing the disutility of work, schooling enhances home production, but the high return to education on earnings is sufficient to generate a positive schooling-participation profile. The earnings profiles in the full-time employment sector are more concave than those found in other structural models of female labor force participation (Eckstein and Wolpin 1989*a*; Van der Klaauw 1996): this is perhaps a result of the endogeneity of the fertility decisions. We find a clear relationship between earnings ability and preference for children. Women with a comparative advantage in market work (or, alternatively, whose earnings profiles are the highest) are those with the lowest marginal utility of children; conversely, women with a strong preference for children are those with lower earnings profiles. Because we estimate a joint model of fertility and labor force participation, we can quantify the lifetime utility effect of a work interruption when a child is born. For a woman who withdraws from full-time employment for 5 years, the utility loss is about 1.5% per year over a 20-year period, provided that she reenters the full-time employment sector after interruption. Shorter withdrawals lead to smaller utility losses. But part-time work during childbirth and child-rearing years does not contribute to a substantial increase in lifetime utility as compared to work interruption.

Several extensions of the model would be desirable. First, imperfect fertility control could be introduced (Montgomery 1992). This would explicitly treat the fertility decisions as uncertain by specifying a birth (or conception) probability structure similar to the job offer probability structure specified in Sauer (1998). Second, incorporating the age distribution of children, as in Hotz and Miller (1988) and Ahn (1995), would help better characterize the household budget constraint in general, and the varying cost of child maintenance in particular. Third, using data from another cohort of women might reveal differences in fertility and labor force participation behavior while holding the constraints, the preference structure and ability constant. Finally, combining this model and a model with endogenous marital status as in Van der Klaauw (1996) into a unified framework would enormously enhance our understanding of the family and work dynamics and of the interaction of economic processes in the market and nonmarket sectors.

Appendix

Table A1
Husband's Log Earnings Equation Estimates

	Always-Married		Ever-Married	
Variable (Z)	Coefficient (ϕ)	SE	Coefficient (ϕ)	SE
Age	.337	.011	.163	.007
Age squared	−.004	.0002	−.001	.0001
Age at marriage	−.328	.065	.101	.010
Age at marriage squared	.005	.001	.001	.0002
Schooling	.190	.041	.143	.016
Schooling squared	−.006	.001	−.006	.001
Constant	5.684	.854	7.812	.143
Adjusted R^2	.4836		.4470	
Person-year observations	7,869		11,788	

NOTE.—Regressions are performed on person-year observations with positive husbands' earnings. Regressions also include eight schooling-age marriage interaction dummies (base category = women with less than 12 years of schooling and age 21 or less at marriage).

Table A2
Ordinary Least Squares Estimates of Wife's Log Earnings Equation

	Always-Married		Ever-Married	
Variable	Part-Time	Full-Time	Part-Time	Full-Time
Schooling	.079	.093	.071	.092
	(.014)	(.008)	(.010)	(.007)
Part-time experience	−.058	−.073	−.076	−.075
	(.020)	(.012)	(.017)	(.010)
Part-time experience squared	.006	.006	.007	.005
	(.002)	(.001)	(.001)	(.001)
Full-time experience	.026	.078	.025	.077
	(.019)	(.007)	(.013)	(.006)
Full-time experience squared	.001	−.002	.002	−.002
	(.001)	(.0003)	(.001)	(.0003)
Constant	7.584	7.998	7.747	8.053
	(.196)	(.115)	(.139)	(.092)
Adjusted R^2	.0858	.2410	.0855	.2340
Person-year observations	1,615	3,268	2,617	4,689

NOTE.—Robust standard errors are given in parentheses.

Table A3
Multinomial Logit Estimates of Labor Force Participation

	Always-Married		Ever-Married	
Variable	Part-Time	Full-Time	Part-Time	Full-Time
Age at marriage	.048	−.020	.016	−.011
	(.018)	(.024)	(.015)	(.018)
Schooling	.128	.234	.045	.176
	(.024)	(.029)	(.022)	(.026)
Part-time experience	.724	.020	.710	−.061
	(.041)	(.046)	(.041)	(.038)
Part-time experience squared	−.029	.007	−.033	.012
	(.004)	(.004)	(.004)	(.003)
Full-time experience	−.013	.522	−.096	.425
	(.027)	(.027)	(.024)	(.021)
Full-time experience squared	.003	−.015	.007	−.012
	(.002)	(.002)	(.002)	(.001)
Husband's real annual earnings/1,000	−.078	−.116	−.033	−.060
	(.009)	(.009)	(.007)	(.006)
Total number of children aged:				
Less than 3	−.328	−1.255	−.321	−1.274
	(.070)	(.104)	(.059)	(.083)
3–6	−.369	−.762	−.504	−.948
	(.080)	(.097)	(.064)	(.075)
6 or more	−.141	.179	−.208	.092
	(.060)	(.068)	(.055)	(.052)
Constant	−3.264	−1.729	−2.092	−1.998
	(.457)	(.546)	(.375)	(.388)
Log likelihood	−11,601.9		−18,794.2	
Person-year observations	15,162		22,701	

Note.—Robust standard errors are given in parentheses. Nonwork is the base category.

Table A4
Actual and Predicted Labor Force Transition Matrices: Always-Married

	Labor Force State (t)			χ^2 Test of Equality (Row)
Labor Force State ($t-1$)	OLF	PT	FT	
OLF:				
Actual	5,389	628	358	
Predicted	5,413	647	341	.697
PT:				
Actual	666	1,521	448	
Predicted	632	1,493	425	.363
FT:				
Actual	380	483	4,524	
Predicted	358	461	4,627	2.001
χ^2 test of equality (column)	1.520	.872	1.809	

Note.—OLF = out of the labor force (nonwork); PT = part-time work; FT = full-time work. There are no statistically significant observations.

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