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The effect of child care and part time opportunities on participation and fertility decisions in Italy

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Abstract. Economic models of household behavior typically yield the prediction that increases in schooling levels and wage rates of married women lead to increases in their labor supply and reductions in fertility. In Italy, low labor market participation rates of married women are observed together with low birth rates. Our explanation involves the Italian institutional structure, particularly as reflected in rigidities and imperfections in the labor market and characteristics of the publicly-funded child care system. These rigidities tend to simultaneously increase the costs of having children and to discourage the labor market participation of married women. We analyze a model of labor supply and fertility, using panel data. The empirical results show that the availability of child care and part time work increase *both* the probability of working and having a child.

JEL classification: J2, C3, D1

Key words: Labor market decisions, fertility, child care

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1. Introduction

The costs of children consist of expenditures for market goods and the opportunity costs of the time spent on child care (see, e.g., Becker 1981; Cigno

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Table 1. Participation and fertility rates 1997

Country	LFP	TFR
Italy	0.41	1.19
France	0.56	1.81
Greece	0.36	1.30
Spain	0.40	1.22
Sweden	0.81	1.92
Denmark	0.82	1.93
UK	0.65	1.81

Sources: OECD (Labour Force Statistics) and Eurostat, 1998

1991; Joshi 1990). Assuming that children are not inferior goods, standard microeconomic theory predicts that the demand for quantity as well as quality will increase with the household's non-labor income. When husbands are not involved in child care activities [as time-providers], an increase in the husband's wage has the same positive effect on the demand for the quantity and quality of children as an increase in non-labor income. Because the wife typically supplies substantial amounts of time to child care activities, an increase in her wage increases household demand for child quantity and quality though an income effect but decreases in demand result from increases in her time cost of child care. One of the main implications of microeconomic models of time allocation in the household is that increases in the wage rates of women should lead to increasing female labor market participation rates and decreasing fertility rates.

In almost all industrialized countries a rise in female participation and a decline in fertility rates have been observed in the last few decades. In Italy, however, fertility has declined dramatically while increases in labor market participation rates have been modest. Table 1 shows that in Italy (as in Spain and Greece) fertility *and* participation rates are substantially lower than in other non-Mediterranean countries.

Our proposed explanations for this apparent anomaly involve the characteristics of the Italian institutional environment, most importantly the particular rigidities and imperfections that are pervasive in the labor market and peculiar features of the publicly-funded child care system. In regards to the labor market, the work rules and wage-policies implemented during the seventies and eighties have served to increase job security for full-time labor market participants, but this benefit for some has come at the cost of a lower probabilities of finding work for new entrants and/or individuals looking for temporary or part-time employment. The fact that part-time employment is extremely rare in Italy is an important factor in accounting for the low employment rates of married women, particularly those with children. As a consequence, married women are forced to choose between no work or full-time work, neither of which is necessarily their preferred option. Married women who choose to work tend to have full-time work commitments, which is not conducive to having large numbers of children. Moreover, even married women who do not work tend to limit their family size, at least in part due to the characteristics of the labor market. Because entry level positions are so hard to find, many children live at home until they find their first "stable" employment. Thus the labor market indirectly imposes large fertility costs on families even when the mother does not work; thus the structure of the Italian labor market both directly and indirectly acts so as to discourage fertility.

The public child care system does not provide services which are of much assistance to married women in terms of reducing the direct costs of participation. In particular, while the quality of publicly-provided child care services is very high in many regions in Italy, there are a limited number of slots available and the hours of child care are typically not compatible with full-time jobs hours. Public child care institutions were developed in an era when the wife either did not participate in the labor market and was responsible for organizing all family activities in a very bureaucratic society, or worked in public sector jobs which required limited time commitments each day [e.g., teaching or public administration]. These institutions are evolving slowly over time and continue to do little to increase the attractiveness of full-time work for women with children.

In light of these considerations, we will consider other factors in the determination of labor participation and fertility besides traditional individual-specific determinants of prices and income. We analyze the effects of several aspects of institutional characteristics using panel data from the Survey on Household Income and Wealth of the Bank of Italy Survey (1991–1995).

In Sect. 2 we discuss several types of market rigidities in Italy which affect the cost of children. Section 3 describes previous empirical literature on the determinants of participation and fertility decisions. In Sect. 4 we outline a model of fertility and participation decisions which takes into account the characteristics of child care and and the availability of part-time work in each individual's region. Section 5 describes the econometric model we will use in the empirical analysis. Section 6 provides a description of the sample used in the empirical analysis and the variables used. Section 7 contains a discussion of the empirical results, and Sect. 8 provides some closing remarks.

2. Institutional rigidities in Italy

In spite of recent institutional changes, the Italian labor market still remains a highly regulated one. Strict rules apply regarding the hiring and firing of workers and permissible types of employment arrangements. The hiring system and the high entry wage as well as very strict firing rules severely restrict employment opportunities for labor market entrants. These labor market regulations have been largely responsible for the high unemployment rates of women and youth (Bertola et al. 1999). The Italian unemployment rate is the highest among industrialized countries, especially the long-term unemployment rate. The unemployment rate of women is twice as high as the male rate (16.8 against 9.5, while the long-term unemployment rates are 11.5% and 6.5% respectively (ISTAT 1998). Due to the high unemployment rates, women may find it hard to take breaks in their working life during childbearing years, because it is difficult to re-enter the labour market (Bettio and Villa 1998).

An important aspect of the rigidity of the labor market is the limited menu of available employment arrangements. Progression towards a more flexibile working hours system has started later in Italy than in other countries and has been much slower. On one hand, unions have traditionally opposed part-time employment fearing that potential divisions of the work force (in terms of working arrangements, demographic characteristics, etc.) could reduce

Table 2. Part-time and women's employment in service

Country	% Part-time	% Women in service sector
Italy	10.1	56
France	12.0	59
Spain	7.0	52
Greece	6.1	54
Denmark	22.8	68
Sweden	24.0	69
UK	25.0	62

Eurostat 1998 % part time among working women.

Table 3. Children < 3 yrs and 3–5 yrs in public child care

Country	% Child care < 3	% Child care 3–6
Italy	6	91
France	23	99
Greece	3	70
Spain	2	84
Denmark	48	82
Sweden	33	72
UK	2	60

Source: Network EEC Child Care 1997

workers' cohesion. On the other hand, under current regulations social contributions paid by employers are strictly proportional to the number of employees, not their hours worked, which makes the employment of two part-time workers more costly than one full-time employee. Moreover the service sector, where part-time work is traditionally more widespread, has not developed as quickly in Italy as in other countries. Table 2 shows the low percentage of part-time workers and the service sector employment share relative to other European countries.

Another source of rigidity is from the Italian child care sector. Child care services are typically inexpensive, relative to private sector alternatives,[1] though their capacity, in terms of number of children and hours per child is extremely limited. The system is highly subsidized but characterized by extreme rigidity in the number of weekly hours available. This makes the service compatible with part-time work but *not* with full-time activities. Having school age children does not necessarily increase the attractiveness of full-time employment since school days often end in mid-afternoon, thus making child care necessary for late afternoon and early evening.[2]

However, a remarkable difference exists between the availability of child care for children under 3 years of age and for children between 3 and 6. Table 3 shows that in Italy the percentage of children less than 3 who are in child care is quite small (6%), while the proportion of children older than 3 in child care is relatively high (91%).

The data reported in this table seems to indicate a correlation between the availability of public child care and women is labor force participation.[3] Table 4 shows that the participation rate of mothers with children under 3 is

Table 4. Mothers' participation by age of the child

Country	Mothers' participation	
	<3 yrs	3–6 yrs
Italy	49	51
France	61	69
Spain	40	44
Greece	47	49
Denmark	84	90
Sweden	86	91
UK	47	62

Source: Eurostat 1998

lower than the participation rate of mothers with children between 3 and 6 years of age (and is certainly lower than is seen in Denmark, Sweden and France)

A main difference concerns the costs of public child care for children before and after 3 years of age, While the first is quite expensive, the second is highly subsidized. Public child care monthly costs for children less than 3 are almost twice as much than costs for children in nursery school, according to the data of special section on public and private services contained in the 1993 survey of the Bank of Italy (about 460 euros for children less than 3 against 255 euros for children older than 3).

While the availability of child care for children older than three is quite uniform across regions, this is not the case for children under three. There are marked differences across regions. The proportion of children less than three years of age in public child care is almost 30% in some areas of the North and only 1–2% in most Southern areas (this ratio is the number of places available divided by the population 0–3 years of age). In the Northern areas the labor market participation rate is about 42%, while in the Southern regions it is about 23% (ISTAT 1998). While private-sector child care services are not the focus of our analysis, we should mention that both public and private child care services are much more prevalent in the North of Italy (Chiuri 2000).

The rigidity and limitations of the supply of publicly-provided child care are somewhat compensated for by a substantial family support system. The number of children under 3 under grandparents' care is 45.7% in households where the mother works and 16.9% in households where the mother does *not* work. Among children between 3 and 6 years of age, the proportion of children under grandparents care is still very high: 39.9% when the mother works and 13.6% when the mother does not work (Indagine Multiscopo *ISTAT* 1998)

The descriptive data presented in this section seem to point to the importance of various market limitations in Italy that are likely to be responsible for the high direct and indirect costs of raising children.

3. Previous empirical results

In the last few years there has been an increasing interest in the effect of institutions (social policies and labor market regulations) on labor market decisions. It has been noted that Italy shares with some other European countries the characteristics of the so-called "Southern model": the lowest level of social protection (especially social expenditures for families and children) and the strictest employment regulations, which together require the family to provide essential "social" services (Bettio and Villa 1998; Ferrera 1996; Saraceno 2000). Cross-country data show that where public support for children is the lowest, women's participation rates are also the lowest (Bradshaw 1997; Gornick et al. 1997).

Analyzing in-kind transfers, it has been shown that the availability of child care services significantly affects women's choices for non-market time versus time spent in paid work. Improvements in child care options as well as variations in their costs have been associated with significant increases in the labor supply of mothers in most countries (Ermisch 1989; Gustaffson 1994, 1995).

Del Boca (1993) and Chiuri (2000) have analyzed the effect of child care on participation of married women and the specific characteristics of the supply of public and private child care systems in Italy. Data on household labor supply and child care use show a strong interdependence between the time use of households members as substitutes for the lack of flexibility and the scarcity of public-provided child care services. Using different data sets the studies arrive to similar conclusions, that is household labour supply depends on child care rationing rather than its costs.

The estimation of the relationship between child care costs and labor supply shows that a reduction in child care costs increases *only* the probability of mothers' part-time employment, but has *no* effect on the probability of working full time (Del Boca 1993). These results raise some concerns, given that part-time employment opportunities are in such a short supply in the Italian labor market.

Empirical studies employing cross-country data have found a high correlation between the proportion of part time jobs and the participation rates of women, in particular married women with children (Meulders and Plasman 1994). The low proportion of part-time does not seem to be coherent with self-reported preferences. A large number of Italian women who are unemployed or out of the labor force report that they would actually prefer to work part-time: surveys at different points of time and different areas of the country have reported similar results. Even among workers more people would like to work fewer paid hours than would like to work more hours at the given hourly wage (European Economy 1995).

Other studies have analyzed the various ways in which the extended family acts as a substitute for the lack of market opportunities. Family networks seems to compensate for the lack of flexibility of the service system. For example, extended family members, most often grandparents, very often provide child care services which complement the limited services provided by publicly-funded day care facilities. Financial support as well as potential help in child care has been shown to significantly increase the probability of the mother's working, and especially has an important effect on the probability of mothers' working full-time (Del Boca 1997).

The role of the family in support of children often extends far beyond the completion of schooling by the children. Because of the limited access to credit and housing markets to individuals without stable employment, the Italian family traditionally provides income support to its children during their usually lengthy search for a stable, "protected" job. This support includes both direct monetary transfers, as well as the provision of housing and other necessities (Cigno et al. 1998; Martinez-Granado and Ruitz Castillo 1998). Imperfections in the Italian credit market such as strict limitations on the size and duration of mortgages, have resulted in parents largely assuming responsibility for providing loans for housing purchases.

In addition to financial transfers from parents to their adult children, parents also provide support by having their mature children live in their home (Giannelli and Monfardini 1998). The proportion of Italian children in the age group of 20–29 living with their parents is more than 70% while in other European countries such as France, Germany, UK it is about 30–35% (Eurostat 1997).

4. A simple model of the fertility-participation decision

In our empirical analysis we will be concerned with determining the relationship between certain "environmental" characteristics, namely the availability of publicly-provided child care and part-time employment opportunities, and fertility and participation decisions. In this section we develop a simple random choice framework in which individual choices are the outcomes of individual preferences and characteristics as well as the nature of the "local" environment in which they live. This framework is used to aid in the specification and interpretation of the econometric model and empirical results reported below. As we will see, the model is not developed and restricted so as to deliver unambiguous comparative statics results. On the contrary, we will see that even a simple model like the one developed here may produce ambiguous comparative statics predictions regarding the effect of environmental variables on household choices.

It is especially difficult to construct a simple framework in which part-time employment availability influences both fertility and participation decisions, primarily because a partial or general equilibrium model of the labor market is required to perform a serious analysis of such issues. That is, one should construct a model in which characteristics of individuals, especially including their preferences and nonlabor incomes, and characteristics of firms, essentially their production technologies and output markets, jointly determine an equilibrium distribution of wage and hours offers. The environment described in our example should be thought of as reflecting the outcome of such an equilibrating mechanism, and therefore may be expected to change if preferences or technologies were altered in some significant manner.

In the Italian labor market there is little difference between the wages (on an hourly basis) of part-time and full-time employees, largely due to institutional constraints. Therefore, we assume that the wage paid w per unit time worked is independent of the number of hours worked, however it will generally be a function of their stock of human capital. Furthermore, we assume that in terms of labor market participation "levels," three are possible. An individual can possibly be observed to (1) not participate; (2) work part-time; or (3) work full-time. For the moment, let us condition on the number of children in the household, n. Let the household utility function be given by $U(x, k)$, where x is parents' consumption, k is a measure of child services, and $k = k(s, n; \theta)$, where s is the time spent by the mother in child-rearing activities and θ is a vector of parameters which completely describes the production technology. Assume that $\partial^2 k/\partial s \partial n > 0$ for all θ, so that the marginal benefit of mother's time in child services is an increasing function of the number of children in the household, an assumption that seems reasonable.

Let the total amount of income available to the household not related to the earnings of the woman be denoted by Y; this includes household nonlabor income and the earnings of husband. There is a monetary cost of having children which is proportional to the number of them in the household, so that total household income excluding the earnings of the wife and direct expenditures on the children is given by $Y - \delta n$. Let the time endowment of the mother be T, and let a full time job require a time commitment of h_f while a part-time job requires h_p, where $0 < h_p < h_f < T$.

Conditional on n and Y, household's first rank order the values of the three alternatives available to them. The value of nonparticipation is given by

$$V_0(n, Y; \theta, \delta) = U(Y - \delta n, k(n, T; \theta)),$$

the value of part-time work is given by

$$V_p(n, Y, w, h_p; \theta, \delta) = U(Y - \delta n + wh_p, k(n, T - h_p; \theta)),$$

and the value of full-time work is given by

$$V_f(n, Y, w, h_f; \theta, \delta) = U(Y - \delta n + wh_f, k(n, T - h_f; \theta)).$$

Given any ordering of V_0, V_p, and V_f, the labor market environment affects the realized choice in the following manner. If $V_0 = \max[V_0, V_p, V_f]$, then the individual chooses not to participate. If $V_p = \max[V_0, V_p, V_f]$, the preferred outcome of the household is to have the wife engaged in part-time employment. Finally, if $V_f = \max[V_0, V_p, V_f]$, the household's preferred alternative is to have the wife work full-time.

Especially in a labor market such as Italy's, the desire to have a certain type of employment relation does not translate into being able to immediately locate one. While full-time jobs are difficult to locate, it is safe to assume that even more "rationing" occurs with respect to part-time ones. The probability of either type of job will be a function of the supply of such jobs by firms in the local labor market and the demand for them by individuals. For example, given that an individual searches for a part-time job, we let π_p denote the probability that she will locate one. Thus, if part-time work is the preferred outcome, there is only a probability of π_p that it will be the observed outcome. Assuming the woman searches for part-time work and doesn't find it, she must restrict her choices to full-time employment and nonparticipation. If $V_f = \max[V_0, V_f]$, she will search for a full-time job. Let π_f denote the probability that a full-time job can be located. Then with probability π_f she will be observed in full-time employment, and with probability $(1 - \pi_f)$ she will be observed in the household's least preferred state, nonparticipation. Of course, if $V_0 = \max[V_0, V_f]$, then she would immediately choose nonparticipation, a choice that by definition is always available to each household.

In the Table below, we present the probabilities associated with each of the outcomes as a function of the preference ordering of the household. In each row of the table we give one of the six possible preference orderings (we ignore the possibility of ties), and the probabilities associated with each observed outcome. Since neither access to full- or part-time jobs is guaranteed, there is a positive probability of being in the nonparticipation state no matter what the preference ordering of the household. On the other hand, we see that three of the six preference orderings lead to part-time employment with a positive probability, and that three orderings may lead to full-time employment. Four of the orderings lead to part-time or full-time employment with a positive probability.

Ordering	Participation state		
	0	p	f
$V_0 > V_p > V_f$	1	0	0
$V_0 > V_f > V_p$	1	0	0
$V_p > V_0 > V_f$	$1 - \pi_p$	π_p	0
$V_p > V_f > V_0$	$(1 - \pi_p)(1 - \pi_f)$	π_p	$(1 - \pi_p)\pi_f$
$V_f > V_0 > V_p$	$1 - \pi_f$	0	π_f
$V_f > V_p > V_0$	$(1 - \pi_p)(1 - \pi_f)$	$(1 - \pi_f)\pi_p$	π_f

Let $p_{ijk}(n, Y)$ denote the probability that a household with n children and income Y have the preference ordering $V_i > V_j > V_k$. Then the probability that the mother is observed to have a part-time job is given by

$$\begin{aligned}\Pr(p; n, Y) &= p_{p0f}\pi_p + p_{pf0}\pi_p + p_{fp0}(1 - \pi_f)\pi_p \\ &= \pi_p(p_{p0f} + p_{pf0} + p_{fp0}(1 - \pi_f)),\end{aligned}$$

and the probability of observing the individual in full-time employment is

$$\begin{aligned}\Pr(f; n, Y) &= p_{f0p}\pi_f + p_{fp0}\pi_f + p_{pf0}(1 - \pi_p)\pi_f \\ &= \pi_f(p_{f0p} + p_{fp0} + p_{pf0}(1 - \pi_p)).\end{aligned}$$

Given the probability distribution over preference orderings, the probability of observing an individual in part-time employment is an increasing function of π_p and a decreasing function of π_f, as seems reasonable. The probability of observing an individual in full-time employment is an increasing function of π_f and a decreasing function of π_p. This is merely to state that, given preferences and (n, y), the probability of observing an individual in a given type of job is an increasing function of the availability of such jobs and a decreasing function of their "competitors." for these positions.

We think of fertility decisions as being made before labor market outcomes are determined, and we will continue to treat as fixed the household income level Y. Given the labor market environment, which is characterized by w, h_p, h_f, π_p, and π_f [we will only think of π_p and π_f as varying across local labor markets], the choice of n essentially amounts to a choice of preference orderings. Formally, the household solves the problem of

$$\max_n EV(n, Y) = \max_n \sum_i \Pr(i; n, Y) V_i(n, y).$$

where $EV(n, Y)$ denotes the expected value associated with income Y and family size n, where the expectation is taken with respect to the conditional probability distribution of labor market states. Written in this manner, it is clear that the choice of n determines both the level of utility realized *ex post* as well as the probability distribution over outcomes.

This framework can also be used to investigate the effect of the availability of publicly-fund child care on the participation and fertility decision. As is the case with any "environmental" variable, the determination of the number of publicly-funded child care spaces in a community can only be adequately investigated using a partial or general equilibrium model. The econometric model we utilize below allows us to estimate consistently the effects of these environmental variables on individual decisions even when the environmental variables are endogenously-determined.[4] For the present, we simply assume that the level of child care availability is predetermined.

We think of the level of child care services as a determinant of the parameters in the child services production technology, $k(s, n; \theta)$. For example, the availability of child care services, particularly those provided at little or no direct cost to the household, may simply reduce the amount of mother's time required to produce a given level of child services, conditional on the number of children in the household. That is, let θ denote the parameter vector characterizing k when a given level of child care services are available, and let θ' be the corresponding vector associated with a higher level of child care services. Then we have $k(s, n; \theta') > k(s, n; \theta)$, and $k(s', n; \theta') = k(s, n; \theta)$ for $s' < s$ and for all n.

The first inequality implies that for the same number of children and parental time input, at least as great a level of child services can be produced in the "better" regions for all levels of s; the second equality states that it is possible to produce the same level of child services (at any family size n) in the "better" regions with a lower level of time input s' $(< s)$.

Without specific restrictions on the forms of the functions k and U it is not possible to say much more about the effects of π_p, π_f, and θ on fertility and participation choices. However, certain reasonable claims might be made given the facts of the Italian labor market. Since part-time employment is the most scarce, it is reasonable to assume that individuals who would prefer part-time employment (as either a first or second choice) are quite likely to be "rationed away" from it. As long as the preferences of individuals are not radically different across local labor markets, we should see that the higher the probability of finding a part-time job, the more likely this state is chosen. While this reasoning might sound a bit circular, recall that (1) we are working with small samples of individuals from each region in Italy; not the entire population of each region (in which case the statement would be tautological), and (2) the proportion of jobs which are part-time is a function of firm behavior in addition to the behavior of individuals on the supply side of the market.

The model also allows us to think about the effect of the labor market environment on fertility decisions in a systematic manner. For example, if women with large families most prefer the part-time option, local labor markets in which part-time work is relatively plentiful will increase the attractiveness of large families (in an *ex ante* sense). Thus characteristics of the local labor market will both partially determine the observed labor market status of married women *conditional* on family size (n), and the (original) choice of family size. Thus the local labor market affects observed labor market status directly through the employment choices available to any given household as well as by shifting household preferences (through the family size effect).

We should note that relationships between environmental characteristics and individual household behavior can be generated by a number of different mechanisms in addition to the one outlined above. For example, the provision of public goods, as reflected in part in the number of child-care slots available in a region, may partly be a function of regional wealth. If wealthy individuals congregate in the same regions, provide high levels of public goods (such as child care services), yet prefer to have the wife out of the market because of these wealth effects, we will mismeasure the availability of child care services on fertility and especially labor market choices to the extent to which we mismeasure household wealth. While systematic sorting of individuals into regions based on household characteristics which are poorly measured can always be claimed to generate what appear to be environmental effects on behavior, the econometric methods we use will tend to minimize the likelihood of estimating spurious contextual effects.

5. The econometric method

In our analysis of fertility and female labor supply, we want to take into account some of the relevant characteristics of the institutional environment, indicators of levels of family support available to the household, and standard demographic characteristics, as well as other factors assumed unobservable to the analyst. One of the limitations of the economic analysis of fertility is the omission of factors such as fecundity, tastes, and other individual and marriage-specific traits which are important factors in explaining the decision to have children. Many, or most, of these individual-specific factors affecting the decision to have a child are unobservable to the researcher. To take into account and isolate these effects we use a fixed-effect model with panel data which is consistent with simple behavioral framework outlined above.

The fixed effects logit estimator allows us to isolate the effects of a subset of the variables included in the analysis on the probabilities of work and fertility allowing for unobserved individual-specific effects which have an *unrestricted* relationship with the included regressors. We use the conditional logit estimator proposed by Chamberlain (1980) to analyse jointly the decisions of having children and working.

The cost of using this rather flexible estimation method is the inability to determine the effect of variables which do not vary over time (at the individual household level) on the probability of having a birth or working in any given period. The conditional maximum likelihood estimators are consistent no matter what the form of the dependence between individual's characteristics and the value of her unobserved "type," and will also be consistent if the "error terms" are correlated across sample members in any manner (Moulton 1990).

When analyzing one binary choice variable, let individual i experience the event in period t with probability given by

$$p(d_{it} = 1 \mid X_{it}, \eta_i) = \frac{\exp(X_{it}\beta + \eta_i)}{1 + \exp(X_{it}\beta + \eta_i)}, \quad i = 1, \dots, N; \quad t = 1, \dots, T;$$

where X_{it} is a vector of covariates associated with individual i in period t, β is an (unknown) associated parameter vector, η_i is an individual-specific, time-invariant error term which is unobservable to the analyst, T is the number of observations available for each household, and there are N households in the sample. The form of the dependence between the scalar random variable η_i and the covariates X_i is not specified; in particular, the estimator for β proposed by Chamberlain is consistent no matter what the form of the conditional distribution of $\eta_i|X_i$. The idea behind the estimator is to find distributions of the data which are functions only of β and not the problematic $\eta_1, \dots, \eta_N$. Define the total number of periods in which the individual experiences the event by $D_i = \sum_{t=1}^{T} d_{it}$. This conditioning method to eliminate the fixed effects can be used for any set D which is greater than 0 and less than T.

In our application, we actually are modeling two decisions simultaneously, the participation decision and the fertility decision. Let d_{it}^j be an indicator variable which takes the value 1 for individual i where $j = f$ for a birth and $j = p$ for labor market participation. We specify the probability that $d_{it}^f = 1$ and $d_{it}^p = 1$ as

$$p(d_{it}^f = 1, d_{it}^p = 1 \mid X_{it}^f, X_{it}^p, \eta_i^f, \eta_i^p) = p(d_{it}^f = 1 \mid X_{it}^f, \eta_i^f) p(d_{it}^p = 1 \mid X_{it}^p, \eta_i^p)$$

$$= \frac{\exp(X_{it}^f \beta_f + \eta_i^f)}{1 + \exp(X_{it}^f \beta_f + \eta_i^f)} \times \frac{\exp(X_{it}^p \beta_p + \eta_i^p)}{1 + \exp(X_{it}^p \beta_p + \eta_i^p)},$$

where X_{it}^j are the exogenous variables in the index function for decision j, β_j is the coefficient vector associated with the exogenous variables X_{it}^j, and η_i^j is the individual specific constant term in the index function for decision j. Just as we do not restrict the form of dependence between X_{it}^j and η_i^j, *we also do not make any assumption concerning the relationship between* η_i^f *and* η_i^p. Given the independence of the decisions f and p conditional on the $X's$ and the $\eta's$, and given that the fixed effects estimator defined below conditions on the $X's$ and eliminates the $\eta's$, the estimator for each decision j is independent of the estimator for the decision j'.

This simple functional form can be used to build likelihood functions which yield consistent maximum likelihood estimators of identified elements of β for each D between 1 and $T - 1$ (see Appendix). In our application of the fixed effects logit estimator, T is at most equal to 3. In this case, subsamples of individuals who experience the event once or twice can be used to estimate β consistently using this method.

Chamberlain proved that the conditional likelihood estimator is consistent and asymptotically normally distributed under standard regularity conditions. In addition to implementing the conditional likelihood estimator on panel data, we also estimate cross-sectional logit specifications. When estimating the logit models, the entire available sample for each year is used. Besides being based on much larger samples, the cross-sectional logit estimator yields coefficient estimates for each variable in the appearing in the index function, even those with values which are time-invariant for each individual, but *if and only if* the condition $\eta_1^f = \eta_2^f = \cdots = \eta_I^f$ and $\eta_1^p = \eta_2^p = \cdots = \eta_I^p$, which is a formal way of stating that households must not systematically differ in terms of any unmeasured variables. We test this restriction using standard "Hausman" tests; if the restriction is rejected, then we conclude that the cross-sectional logit estimates are inconsistent, although the fixed effects estimates will be consistent. If the restriction cannot be rejected, we will have some preference for the cross-sectional logit estimates, which are consistent in this case, since they are based on more sample information (and hence are more efficient) and produce estimates of all coefficients in the model, not only those associated with variables that vary over the sample period.[5]

6. The data

The empirical analysis utilizes a three-year panel from the Bank of Italy's Survey of Households' Income and Wealth (1991–1995). The Bank of Italy survey contains detailed information on the incomes and wealth of family members, several characteristics of the workplace (such as wages and hours of work), and socio-demographic characteristics of the households (age of the members of the family and the number of children). The sample design of the Bank of Italy panel, which is somewhat unorthodox, is well described in detail in Trivellato (1997).

For purposes of our analysis of fertility and labor market participation we have selected sample households with married women in the age range 21–45 so as to exclude those who might be enrolled in school or in retirement or semi-retirement (which occurs at relatively young ages because of the historically generous Italian pension system). For the analysis of fertility, the age restriction serves to ensure that women included in the final sample will have a high probability of being fecund. The sample size, after excluding women who didn't meet the age criteria or who had missing information on the variables included in the analysis, was 1708.

We note that though our discussion has emphasized the distinction between full- and part-time employment options, there are too few women working at part-time jobs to estimate separately equations determing the probabilities of part-time and full-time employment. Only 9% of our sample work part time. Therefore we have folded these two categories into one, and only consider the employment/no employment option. At a minimum, the proportion of part-time jobs in the local labor market can be viewed as an indicator of the flexibility of employment relationships. The attractiveness of employment to married women with household organization and possibly child care responsibilities will be a function of the flexibility of employment relations in the local market.

In order to use the conditional likelihood estimator we need to limit our analysis to the women who change states over the observation period. For the participation analysis, our sample includes wives who worked at least one period and less than three periods (227 women). For the fertility analysis, there were 201 women who had a least one birth and less than three over the three periods.

The dependent variables are whether the wife is working at the time of the interview and whether or not she had a child in the last two years. For each sample member, we have three observations on each of the two dependent variables. Only a few of the independent variables are not time-invariant: We include in our analysis variables related to:

Personal characteristics. Wife's age, family income which is constructed as total family income minus wives' earnings (in euros divided by 1000).

Family support. Variables indicating 1) the transfer the family has received

Table 5. Descriptive statistics of variables (means and standard deviations)

Variables	1991	1993	1995
Fertility	0.099	0.095	0.094
	(0.262)	(0.261)	(0.325)
Participation	48.5	48.0	47.8
	(0.371)	(0.367)	(0.377)
Household income (euros)	21.252	22.345	23,433
Positive values	(1.587)	(1.033)	(1.289)
Age of the wife	34	36	38
	(12.5)	(12.5)	(12.4)
Family transfers (euros)	2,336	1,639	2,632
Positive values	(884)	(667)	(856)
Number of children	1.58	1.63	1.76
	(1.10)	(1.11)	(1.12)
Wife schooling	10.33	10.44	10.43
	(4.40)	(4.5)	(4.40)
Child care	7.0	9.43	9.73
	(7.6)	(7.7)	(7.8)
Parents alive	88.5	87.7	86.2
	(37.6)	(37.7)	(37.7)
Part time	6.0	6.88	6.90
	(4.56)	(4.77)	(4.78)

from relatives during the year of the interview (euros divided by 1000) 2) a dummy variable indicating whether one of the parents is still alive.

In order to measure the impact of rigidities of the aggregate labor market and publicly-provided goods on household decisions we have merged our panel data with regional data on child care facilities and part-time jobs.

Child care system. As an indicator of the characteristics of the child care system, we use the ratio of the number of child care places available (for children under 3 years of age) to the number of children 3 years of age or less by area of residence in 1991, 1993, and 1995.

Labor market. As an indicator of the probability of locating a part-time job, we use the ratio of the number of part-time jobs to total employment in the region.

Table 5 reports descriptive statistics for the variables used in the empirical analysis for the three years we have considered. The evidence from these data are in accordance with the premises of our earlier arguments. Fertility rates (the proportion of women who had a child in each of the three two-year periods) are very low and tend to decline during the period 1991–1995. Participation (proportion of women working) also declines over the period (from 0.48 to 0.47). Family income increases during the period. The amount of transfer income does not change much during the period. The proportion of households in which one of the parents is still alive is 88% and decreases slightly during the period.

We also present the means of the "environmental" data by region (Table 6). It shows that there exists quite a remarkable variability in child care availability across regions, with a far higher supply of facilities in the Northern regions compared with Southern regions. Part-time employment shows much less variability and does not appear to be significantly different in the North and South.

Table 6. Child care and part time by region 1991–1995[1]

Regions	Child care			Part time		
	1991	1993	1995	1991	1993	1995
Piemonte-Val d'Aosta	14	16	16.5	5.8	5.9	5.9
Lombardia	13	13.6	13.8	6.5	5.6	6.8
Trentino	7.3	11.0	11.0	7.1	7.3	7.8
Friuli-Veneto	7.1	8.5	8.6	7.0	6.5	6.8
Liguria	12.0	11.0	10.1	6.5	5.3	6.3
Emilia	29	28.2	28.4	7.3	6.2	6.6
Toscana	11.	11.6	11.7	7.0	6.1	7.1
Umbria	12.6	11.8	11.5	7.0	5.6	6.7
Marche	11.5	13.5	13.2	5.2	5.2	5.9
Lazio	9.9	10.0	10.4	5.1	4.6	5.6
Abruzzo-Molise	4.9	3.5	4.5	4.9	4.6	5.6
Campania	1.0	0.9	1.0	3.9	3.9	4.5
Puglia	6.2	4.9	4.8	4.5	4.2	6.2
Basilicata	5.2	5.8	5.8	5.0	5.5	5.7
Calabria	1.2	1.3	1.2	6.5	7.4	7.3
Sicilia-Sardegna	3.5	4.2	4.0	5.0	5.8	5.8

[1] Annuario Statistico Italiano ISTAT (1997) and Statistiche della Previdenza, della Sanità e Assistenza Sociale ISTAT (1995).

7. Empirical results

Table 7 and 8 report the fixed effect and the logit estimates using pooled cross-sections. In the first column (FE) the conditional logit estimates are reported and in the second column (CS) estimates of the logit specifications on the pooled sample obtained by pooling the cross sections for 1991, 1993, 1995 are presented. Besides being based on much larger samples, the cross-sectional logit estimator yields coefficient estimates for each variable appearing in the index function, even those with values which are time-invariant for each individual.

We can compare the cross-sectional and fixed effects of common parameters when such a comparison is possible, i.e., when the coefficient is associated with a time-varying variable. The estimates of comparable parameters are in some cases substantially different between the cross-sectional logit and conditional likelihood estimators. In general, the cross-section estimates are larger in absolute value and estimated more precisely than are the fixed effect ones (which is to be expected since they are effectively based on much larger sample sizes). We will discuss some of these differences first.

The fixed effect and the cross-sectional logit estimators of the effect of household income on participation are both negative. This is not true in the fertility equation (Table 8), where the fixed effect estimate of household income coefficient is positive (though not significant) while the cross-section estimate is negative.

The other results are quite similar across the two estimation methods. The effects of personal characteristics conform to other findings reported in the recent literature on fertility and women's labor market participation using cross-sectional data (Colombino and Di Tommaso 1996; Del Boca 1997). The wife's age has a negative effect on participation and fertility. Wife's schooling has a

Table 7. Participation equation estimates (Asymptotic standard errors in parentheses)

Variable	FE	CS
Household income	−0.068	−0.105*
	(0.045)	(0.029)
Family transfers	0.052	0.112*
	(0.032)	(0.010)
Age	−0.105	−0.062*
	(0.077)	(0.005)
Child care	0.056*	0.045*
	(0.024)	(0.011)
Part-time	0.065	0.038*
	(0.034)	(0.012)
Schooling	–	0.176
		(0.017)
Parents alive	0.022*	0.048
	(0.010)	(0.023)
1993	–	0.122
		(0.012)
1995		0.345
		(0.176)
Constant		4.906
		(3.236)
Spec. test		109.922

Table 8. Fertility equation estimates (Asymptotic standard errors in parentheses)

Variable	FE	CS
Family income	0.069	−0.056
	(0.051)	(0.027)
Family transfers	0.039	0.052*
	(0.020)	(0.022)
Age	−0.077*	−0.221*
	(0.035)	(0.031)
Child care	0.057*	0.022
	(0.036)	(0.020)
Part-time	0.033	0.031*
	(0.028)	(0.010)
Schooling	–	0.041
		(0.032)
Parents alive	0.049*	0.269*
	(0.020)	(0.110)
1993	–	−0.021
		(0.017)
1995	–	−0.022
		(0.014)
Constant		5.017
		(2.101)
Spec. test		108.963

positive effect on participation and fertility (this coefficient cannot be estimated using the fixed effects estimator). The positive effect of wife's schooling on fertility can be interpreted in part as a permanent income effect, given that father's education is not included in the analysis (assortative mating).

The first variable that we have introduced as indicator of potential family support is the amount of family transfers. This variable has a positive effect on the likelihood of women participating and having children. The estimated effects of transfers agree with the results obtained in other previous studies. The decisions of working and having children are positively affected by parents financial support. Studies of intergenerational transfers have shown that recipients are also more likely to have been denied credit than the rest of the population (50% of family transfer recipients have been denied credit from financial institutions), confirming an important role for the family as a system of household financing (Cigno et al. 1998).

The amount of family transfers is potentially endogenous. For example, it has been found that relatives are more likely to make transfers to families if those families have children present (Mayer and Engelhardt 1994). We found that the estimates of other parameters were relatively insensitive to the omission of this variable. Therefore, while the coefficient estimate associated with family transfers may itself be biased, this potential bias does not seem to affect the other estimates.

We have then analyzed the effect of other indicators of potential family support such as the presence of at least one parent of the wife. We believe that this variable can be interpreted as a potential opportunity for child care (in conditions of limited public child care facilities). Having one parent alive increases both the probabilities of child-bearing and labor market participation, though the effects on fertility are quite a bit larger.

Now consider the effect of environmental characteristics. The fixed effect coefficient estimate of child care availability is positive in both equations, and is at least marginally significant in both. The cross-sectional estimate of the parameter is significant for the participation equation, but this is not the case in the fertility equation. According to our modeling framework, the impact of child care availability on both fertility and participation is predicted to be positive, so that the fixed effects estimates are consistent with this hypothesis.

The fixed effects estimates of the estimate of the part-time are positive in both the fertility and participation equation, but are only significant in the participation equation. This result is consistent with the modeling framework, in the sense that the effect of a flexible labor market was more "direct" in the participation decision than in the fertility decision, although it was expected to be positive in both. The cross section estimates are also positive and significant in both equations. Finally the year dummies for 1993 and 1995 capture the effect of changes in the macroeconomic conditions. The year dummies are positive and marginally significant in the participation equation and negative and non significant in the fertility equation.

Throughout this section we have spent some time interpreting and comparing the fixed effects and cross-sectional estimates of the parameters characterizing the model. While the FE estimator measures only the effect of the variation over the period, the cross-section estimator measures both the effect of the regional variability on the dependent variables at a point in time and the time variation. Which are to be preferred? The answer to this question appears to be that the fixed effects estimates are the only ones in which we should put much faith. The test statistics reported at the bottom of Tables 7 and 8 indicate over-whelming rejection of the null hypothesis (of the equality of all unobservables across households). The rejection of this null implies that the cross-sectional estimate are inconsistent and hence biased. Since the fixed effects estimates are consistent under unspecified forms of heterogeneity, and since they should tend to "purge" the estimates of spuriousness, we are pleased to see that fixed effects estimates of the environmental influences on household behavior are consistent with our the predictions of our modeling framework and reasonably precisely estimated.

8. Elasticities and simulations

The conditional maximum likelihood estimator allows us to estimate consistently a subset of the parameters in β, namely those coefficients associated with variables which change over time for at least a subset of sample members. In particular, if the probability that individual experiences the event in period t is given by

$$p(d_{it} = 1 \mid X_{it}, \beta, \eta_i) = \frac{\exp(X_{it}\beta + \eta_i)}{1 + \exp(X_{it}\beta + \eta_i)},$$

the conditional maximum likelihood allows consistent estimation of the subvector $\tilde{\beta} \subseteq \beta$. Even if all elements of X_{it} vary over time for some or all individuals, so that $\tilde{\beta} = \beta$, the conditional m.l. estimator does not provide consistent estimates of η_i. Using the c.m.l. means that we can never consistently estimate the function $p(d_{it} = 1 \mid X_{it}, \beta, \eta_i)$, so that we cannot compute elasticity measures built around this function, such as the elasticity of the probability of employment with respect to child care availability.

Even though the c.m.l. estimator of $\tilde{\beta}$ does not allow us to perform many of the comparative statics exercise we would like to, it is possible to conduct certain "experiments" which are of substantive interest. Since the c.m.l. estimator works off the relationship between the timing of the dependent events and intertemporal co-movements in the exogenous variables, the types of experiments we can conduct will revolve around the timing of events rather than the number of them. For example, given consistent estimators of $\tilde{\beta}$, we can consider the following experiment. Say that an individual experiences the event in one of two periods, so that either $d_{i1} = 1$ and $d_{i2} = 0$ or $d_{i1} = 0$ and $d_{i2} = 1$. Start from the point in which all elements of X_{it} are constant over time, so that $X_{it} = X_i$ for all t. Then at the baseline, the conditional probability that the individual works in period 2 given that she works in exactly one period is $p(d_{i1} = 0, d_{i2} = 1 \mid X_i, \eta_i, D_i = 1) = 0.5$. Say that we increase the second period value of the k^{th} variable from X_i^k to $(1+\delta)X_i^k$, so that $\hat{X}_{i1}^k = X_i^k$ and $\hat{X}_{i2}^k = (1+\delta)X_i^k$, where δ is a "small" positive number. Now if all the variables except X_{it}^k are constant, then

$$\begin{aligned} p(d_{i1} = 0, d_{i2} = 1 \mid X_{i1}, X_{i2}, \eta_i, D_i = 1) &= \frac{\exp(X_{i2}\beta)}{\exp(X_{i1}\beta) + \exp(X_{i2}\beta)} \\ &= \frac{\exp((X_{i2} - X_{i1})\beta)}{1 + \exp((X_{i2} - X_{i1})\beta)} \\ &= \frac{\exp(\delta X_i^k \beta^k)}{1 + \exp(\delta X_i^k \beta^k)}. \end{aligned}$$

Now as long as $\beta^k \in \tilde{\beta}$, so that it can be consistently estimated using the c.m.l.

Table 9. Elasticities of child care, part time and family support

Variables	Participation	Fertility
Child care	0.296	0.198
Part time	0.244	0.124
Family transfers	0.125	0.079
Parents alive	0.022	0.854
Non labor income	0.197	0.168

estimator, we can evaluate the effect of increasing the value of the k^{th} regressor by the proportion δ between periods one and two on the conditional probability of experiencing the event in the second period. The change in the probabilities is $\frac{\exp(\delta X_i^k \beta^k)}{1+\exp(\delta X_i^k \beta^k)} - 0.5$, and the proportionate change in the value of X^k is δ. Therefore we can define the elasticity

$$E_\delta^k = \frac{\left[\frac{\exp(\delta X_i^k \beta^k)}{1+\exp(\delta X_i^k \beta^k)} - 0.5\right] \Big/ 0.5}{\delta X_i^k / X_i^k}$$

$$= \frac{\frac{2 \times \exp(\delta X_i^k \beta^k)}{1+\exp(\delta X_i^k \beta^k)} - 1}{\delta}.$$

Using a value of $\delta = 0.1$ and the sample average values of the time-varying X_{it}, we compute $E_{0.1}^k$ for all regressors k which have coefficients in $\tilde{\beta}$ [i.e., that have coefficients which are estimable using the c.m.l. estimator]. These elasticities allow us to assess the importance of the changes in the time path of the regressor in question on the timing of events, but not on the number of them. For example, consider the variable child care availability. Beginning from a time-invariant environment [i.e., one in which all regressors are fixed over time], for an individual who works in one of two periods, $E_{0.1}^k$ is the ratio of the percentage change in the probability of working in the second of the two periods with respect to a proportionate increase of 0.1 child care availability in period 2 with respect to period 1. We cannot address how the probability of working in either of the two periods responds to a proportionate change of 0.1 in the second period value of child care avialability. Table 8 shows the elasticities obtained from our estimates derived according to the method described above.

Using the parameters obtained in our estimates we compute elasticities of the time-varying factors, the regional variables, child care and part-time, the family support variables (parents alive and family transfers), and family income. For example, increasing child care availability by 10% in the second period with respect to its first period level increases the relative odds of working in the second period to the first by 0.296, and changes the relative odds of having a child in the second period to having one in the first period by 0.198. In general, the elasticity estimates are not large, indicating that the responsiveness of life cycle decisions to changes in the timing of these exogenous variables is modest. The one exception is the elasticity of the timing of births with respect to the presence of parents. Aside from this elasticity, it is interesting to note that the largest elasticity estimates correspond to environmental varibles, not variables characterizing the household's characteristics.

9. Conclusions

In this paper we have argued that several institutional rigidities are among the factors explaining the low fertility and low participation rates observed in Italy. The limited availability of part-time employment and the limited availability of affordable child care services increase the costs of working for mothers, making it difficult to participate in the labor market without other relatives' support.

To capture the impact of "environmental" variables on household behavior, we have considered a simple model of labor supply and fertility in which rationing and market imperfections can be introduced in a simple but intuitively-appealling way. In estimating the model, we included several variables reflecting levels of potential and actual family support as well as institutional characteristics of the regional child care system and local labor market in order to explicitly take into account relevant constraints that Italian households face when making their labor market and fertility decisions. Even if some of elasticities are not large, our results seem to indicate that labor force participation and fertility decisions are both affected by similar forces. The decisions to work and have a child are positively influenced by the available supply of public child care as well as the availability of part time jobs. The empirical results also indicate that the availability of family support, both in the form of transfers and in the form of the presence of parents, increases *both* the probability of market work and having children.

In some sense, it appears that Italy is stuck in a "low female participation rate" equilibrium in which one of the major reasons for low participation rates is the mismatch between the types of jobs sought by married women with children and the types of jobs offered (full-time). It would appear that this imbalance could be addressed by increasing the provision of child care, which would simultaneously increase job opportunities for women and reduce the costs of taking full-time jobs. It may well be the case that increasing the provision of public child care (in terms of number of slots and hours provided per day) could be financed by slight increases in payroll taxes. Given the relatively underdeveloped private child care system, such a change could be welfare-enhancing. Our empirical results also indicate that by increasing the flexibility of employment relationships, more women would find it attractive to enter the market. Of course, to analyze the welfare effects of such a change one requires knowledge of the implications for wages and the employment status of other household members.

Appendix 1

Let individual i experience the event in period t with probability given by

$$p(d_{it} = 1 \mid X_{it}, \eta_i) = \frac{\exp(X_{it}\beta + \eta_i)}{1 + \exp(X_{it}\beta + \eta_i)}, \quad i = 1, \ldots, N; \quad t = 1, \ldots, T;$$

where X_{it} is a $(1 \times k)$ vector of covariates associated with individual i in period t, β is an (unknown) associated $(k \times 1)$ parameter vector, and η_i is an individual-specific, time-invariant error term which is unobservable to the analyst. Since the probability that an individual experiences the event in any period t, conditional on X_{it} and η_i is independent of the probability that she experiences the event in any other combination of periods, the probability of any given sequence $d_{i1}, \dots, d_{iT}$ given $X_i \equiv (X_{i1}, \dots, X_{iT})$ and η_i is

$$p(d_{i1}, \dots, d_{iT} \mid X_i, \eta_i) = \frac{\prod_{t=1}^{T} \exp[d_{it}(X_{it}\beta + \eta_i)]}{F_i}, \tag{9.1}$$

where $F_i = \prod_{t=1}^{T}[1 + \exp(X_{it}\beta + \eta_i)]$.

The form of the dependence between the scalar random variable η_i and the covariates X_i is not specified. The estimator for β proposed by Chamberlain is consistent no matter what the form of the conditional distribution of $\eta_i|X_i$. The idea behind the estimator is to find distributions of the data which are functions only of β and not the problematic $\eta_1, \dots, \eta_N$. Define the total number of periods in which the individual experiences the event by $D_i = \sum_{t=1}^{T} d_{it}$. First, consider the event $D_i = 1$. The probability that $D_i = 1$ is given by

$$p(D_i = 1 \mid X_i, \eta_i) = F_i^{-1}[\exp(X_{i1}\beta + \eta_i) + \cdots + \exp(X_{iT}\beta + \eta_i)].$$

This expression is the probability that the individual experiences the event in period one but not in the other periods plus the probability that the individual experiences the event in period two but not in the other periods, and so on. Now given that $D_i = 1$, the conditional probability that the individual experiences the event in period t is

$$\begin{aligned}
&p(d_{it} = 1, d_{is} = 0, \forall s \neq t \mid D_i = 1, X_{it}, \eta_i) \\
&\quad = \frac{p(d_{it} = 1 \mid X_{it}, \eta_i) \prod_{s \neq t} p(d_{is} = 0 \mid X_{is}, \eta_i)}{p(D_i = 1 \mid X_i, \eta_i)} \\
&\quad = \frac{\dfrac{\exp(X_{it}\beta + \eta_i)}{F_i}}{\dfrac{\sum_{s=1}^{T} \exp(X_{i3}\beta + \eta_i)}{F_i}} \\
&\quad = \frac{\exp(\eta_i)\exp(X_{it}\beta)}{\exp(\eta_i)\sum_{s=1}^{T} \exp(X_{is}\beta)} \\
&\quad = \frac{1}{1 + \sum_{s \neq t} \exp((X_{is} - X_{it})\beta))}.
\end{aligned}$$

Now consider the case for which $D_i = 2$, assuming that $T > 2$. The probability that the individual experiences the event in periods t and periods t' but not in any other period is given by

$$\begin{aligned}
&p(d_{it} = 1, d_{it'} = 1, d_{is} = 0 \ \forall s \neq t, t' \mid X_i, \eta_i) \\
&\quad = F_i^{-1}[\exp(X_{it}\beta + \eta_i) \times \exp(X_{it'}\beta + \eta_i)].
\end{aligned}$$

The probability that individual i experiences the event in exactly two of the periods is given by

$$p(D_i = 2 \mid X_i, \eta_i) = F_i^{-1}\left[\sum_{j=1}^{T}\sum_{k>j}^{T}[\exp(X_{ij}\beta + \eta_i) \times \exp(X_{ik}\beta + \eta_i)\right],$$

which is the sum of the probabilities of the $T(T-1)/2$ ways in which the event can occur twice in T periods. Then the conditional probability that the individual experienced the event in period t and period t' given that she experienced the event twice in T periods is

$$\begin{aligned}
&p(d_{it} = 1, d_{it'} = 1, d_{is} = 0\ \forall s \neq t, t' \mid D_i = 2, X_i, \eta_i) \\
&\quad = \frac{\dfrac{\exp(X_{it}\beta + \eta_i)\exp(X_{it'}\beta + \eta_i)}{F_i}}{\dfrac{\sum_{j=1}^{T}\sum_{k>j}^{T}[\exp(X_{ij}\beta + \eta_i) \times \exp(X_{ik}\beta + \eta_i)]}{F_i}} \\
&\quad = \frac{\exp(2\eta_i)\exp(X_{it}\beta)\exp(X_{it'}\beta)}{\exp(2\eta_i)\sum_{j=1}^{T}\sum_{k>j}^{T}[\exp(X_{ij}\beta) \times \exp(X_{ik}\beta)]} \\
&\quad = \frac{\exp([X_{it} + X_{it'}]\beta)}{\sum_{j=1}^{T}\sum_{k>j}^{T}\exp([X_{ij} + X_{ik}]\beta)} \\
&\quad = \frac{1}{\sum_{j=1}^{T}\sum_{k>j}^{T}\exp([(X_{ij} + X_{ik}) - (X_{it} + X_{it'})]\beta)}.
\end{aligned}$$

This conditioning method to eliminate the fixed effects can be used for any set D which is greater than 0 and less than T. In particular, let $D_i = k$, $1 \leq k < T$, and let $E_i = (e_{i1}, \ldots, e_{ik})$, where the $\{e_i\}$ denote the k time periods in which individual i experiences the event. Then we have that

$$\begin{aligned}
&p(d_{ie_{i1}} = 1, \ldots, d_{ie_{ik}} = 1, d_{is} = 0, s \notin E_i \mid D_i = k, X_i, \eta_i) \\
&= \frac{1}{\exp(\sum_{j_1=1}^{T-k}\sum_{j_2>j_1}^{T-(k-1)}\ldots\sum_{j_k>j_k-1}^{T}\{(X_{ij_1} + X_{ij_2} + \cdots + X_{ij_k}) - \sum_{t\in E_i} X_{it}\}\beta)}
\end{aligned} \tag{9.2}$$

In our application, we actually are modeling two decisions simultaneously, the participation decision and the fertility decision. Let d_{it}^{j} be an indicator variable which takes the value 1 for individual i in period t if event j is observed, where $j = f$ for a birth and $j = p$ for labor market participation. We specify the probability that $d_{it}^{f} = 1$ and $d_{it}^{p} = 1$ as

$$\begin{aligned}
&p(d_{it}^{f} = 1, d_{it}^{p} = 1 \mid X_{it}^{f}, X_{it}^{p}, \eta_i^{f}, \eta_i^{p}) = p(d_{it}^{f} = 1 \mid X_{it}^{f}, \eta_i^{f})p(d_{it}^{p} = 1 \mid X_{it}^{p}, \eta_i^{p}) \\
&\quad = \frac{\exp(X_{it}^{f}\beta_f + \eta_i^{f})}{1 + \exp(X_{it}^{f}\beta_f + \eta_i^{f})} \times \frac{\exp(X_{it}^{p}\beta_p + \eta_i^{p})}{1 + \exp(X_{it}^{p}\beta_p + \eta_i^{p})},
\end{aligned}$$

where X_{it}^j are the exogenous variables in the index function for decision j, β_j is the coefficient vector associated with the exogenous variables X_{it}^j, and η_i^j is the individual specific constant term in the index function for decision j. Just as we do not restrict the form of dependence between X_{it}^j and η_i^j, *we also do not make any assumption concerning the relationship between* η_i^f *and* η_i^p. Given the independence of the decisions f and p conditional on the $X's$ and the $\eta's$, and given that the fixed effects estimator defined below conditions on the $X's$ and eliminates the $\eta's$, the estimator for each decision j is independent of the estimator for the decision j'.

Thus we are able to consistently estimate β_j using only the information on the outcomes d^j and the X^j, even though the probabilistic model allows for relatively general forms of dependence between the fertility and the participation decision. The brief discussion of the fixed effects estimator considers the univariate choice problem without any loss of generality.

This simple functional form can be used to build likelihood functions which yield consistent maximum likelihood estimators of identified elements of β for each D between 1 and $T-1$. In our application of the fixed effects logit estimator, T is at most equal to 3. In this case, subsamples of individuals who experience the event once or twice can be used to estimate β consistently using this method. Let the subsample for which $D_i = 1$ be denoted by S_1 and let S_2 denote the subset of sample members for which $D_i = 2$. Then we define the conditional maximum likelihood estimator as

$$\hat{\beta}_c = \arg\max_{\beta}\{L_1(\beta) + L_2(\beta)\},$$

where

$$L_1(\beta) = \sum_{i\in S_1}\sum_{t=1}^{3} d_{it}\left\{-\ln\left(1 + \sum_{s\neq t}\exp[(X_{is} - X_{it})\beta]\right)\right\}$$

and

$$L_2(\beta) = \sum_{i\in S_2}\sum_{t=1}^{2}\sum_{t'>t}^{3} d_{it}d_{it'}\left\{-\ln\left(\sum_{j=1}^{2}\sum_{k>j}^{3}\exp([(X_{ij} + X_{ik}) - (X_{it} + X_{it'})]\beta)\right)\right\}.$$

Chamberlain proved that the conditional likelihood estimator is consistent and asymptotically normally distributed under standard regularity conditions.

Endnotes

1 Private child care costs are on average much higher than comparable public ones. According to Bank of Italy data, the monthly costs for public child care for children are much lower than the private ones (on average 350 against 650 euros) Bank of Italy 1993.

2 School hours per week are 27 in Italy, and 30 in Denmark, 35 in France and 33 hours in the UK (Gornick et al. 1997).

3 The UK seems to be the exception if we look at Table 1, but it is not if we look at Table 4 which reports the proportion of mothers with children in the age range 0–3. The proportion for the UK is actually lower than the one in Italy.

[4] The econometric model will only produce consistent parameter estimates under certain assumptions regarding the form the endogeneity takes. We will discuss these issues further in the econometric section.

[5] The asympototic standard error estimates reported for the cross-sectional pooled logits are not "robust" standard errors. Estimates of robust standard errors, which allow for some degree of model misspecification, were computed and were very close to those reported (i.e., no greater or less than 10% for all coefficient estimates). In conducting the Hausman test, it is necessary that the estimator that is inconsistent under that alternative hypothesis (in this case, that fixed effects are present) be efficient under the null. Therefore to carry out the test we must assume that the assumed error structure is the correct one for the data, which implies that it would be inappropriate to use the robust covariance matrix in forming the test statistic. As a practical manner, we should note that the null hypothesis would be decisively rejected no matter which covariance matrix was employed.

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