Are young cohorts of women delaying first birth in Mexico?

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Abstract. In the last decades female permanent sterilisation became the most used method of contraception in Mexico. During this time the demand for pills, condoms and other short-term contraceptives fell consistently. The shift in the demand for contraceptives raises concerns among demographers that the timing of children may remain unchanged regardless of observed reductions in period fertility rates. This paper assesses such ideas in the context of the timing of a first child using duration models as the main analysis tool. Findings suggest that young cohorts of women are effectively delaying first birth relative to the experience of older generations.

JEL classification: J13, J15, C41.

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1 Introduction

In the last twenty years female permanent sterilisation (FPS) has become the most popular method of contraception in Mexico. While in 1978 an eight per cent of users of contraception demanded FPS, in 1998 that figure was estimated to be as high as fifty per cent. During the same period of time the demand for pills, condoms and other short-term contraceptives dropped sharply and consistently (INEGI 2001a). These tendencies in the demand for contraceptives have raised concerns that the timing of children in Mexico might remain unchanged despite the reduction of fertility rates in the last few

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decades. The impression is that women might be cutting off fertility at high parities without modifying significantly their reproductive calendar — adopting no contraception until they reach their desired number of children and then using definite natal control to limit lifetime fertility (see Welti 1997, Gomez 1996, Mier y Teran and Rabell 1990, Zavala de Cosio 1989). Such behaviour, if present, implies a sustained pattern of early entry into motherhood which potentially reduces the chances of Mexican women to accumulate human capital, to improve their welfare, and to promote economic growth in Mexico as a whole. An early entry into motherhood might also increase the overall number of children a woman has over her lifetime and affect the pace of subsequent fertility (for more on this topic see for instance Gustafsson 2001, Chen and Morgan 1991, Heckman et al. 1985, Bloom and Trussell 1984, Bumpass et al. 1978).

The present paper contributes to this discussion by reporting a study on the timing of first birth in Mexico. The main objective is to test whether or not young cohorts of Mexican women are effectively delaying first birth with respect to older generations. To accomplish this objective duration models are estimated using individual-level data from the National Survey of Demographic Dynamics 1997 (INEGI 1999). To the knowledge of the author no previous study has discussed systematically — and with the use of advanced econometric techniques — the issue of first birth postponement in Mexico.

Various econometric aspects of the present work are worth noting. First, unlike the common practice in the analysis of transition data, the hazard function is estimated in a semi-parametrical fashion so that no a priory restrictions on the form of the baseline hazard are imposed. Hence, the possibility of inducing bias due to misspecification of the functional form of the hazard is avoided. Second, as suggested by Heckman and Singer (1984), unobserved individual heterogeneity is controlled for estimating a discrete approximation of the distribution of unobservables. This non-parametric technique reduces potential sensitivity of the results to prior distributional assumptions about unobservables that are otherwise required. Finally, the ‘stayer-mover’ problem induced by the presence of individuals who remain childless until the end of their fertile life is explicitly taken into account avoiding, once again, potential bias due to misspecification.

The remainder of this paper is organised as follows. Section two describes the main stylised facts on population issues in Mexico. Section three presents the data and deals with the definition of all variables used in the analysis. Section four discusses econometric issues and section five presents empirical results. Finally, section six concludes.

2 Stylised Facts and institutional background

Mexico was a low-populated country until the start of the 1930s, at which time its main population expansion began. Between 1930 and 1970 population grew at increasing rates, and for the period 1970-1980 the annual rate of population growth reached a historic maximum of 3.32%. Since then population has consistently grown at declining rates and, according to available data, in year 2000 it was as low as 1.85% (CONAPO 2001). Such reductions in the rate population growth have been fundamentally driven by changes in individual fertility behaviour. In fact, since the end of the 1970s the total fertility rate (TFR) of the county dropped from 6.8 to less than 3 children per woman. This fall in the TFR was accompanied by reductions in all age-specific fertility rates.[1]

Many factors have been associated with the reduction of fertility in Mexico. Infant mortality rate fell in the period 1970-2000 from 68.5 to 17.5 deaths per 1,000 births (INEGI 2000; 2001b). During the same period, average education increased from 3.4 to 7.6 years of schooling, and female participation in the labour force increased from 11 to 26.6 per cent (INEGI 2001b, World Bank 2001). Indigenous and non-indigenous individuals, however, face different realities. For instance, the infant mortality rate among indigenous people in year 2000 was estimated to be as high as 38.5 deaths per 1,000 births, well above the 17.5 deaths per 1,000 births corresponding to Mexico as a whole. Similarly, education among indigenous individuals is estimated to be in average three years lower than education among non-indigenous individuals (INEGI 1999). The reduction in period fertility rates has also been a response to policy actions undertaken by the Mexican government. After a long tradition of pro-natal policy, in 1973 the Mexican government launched an aggressive programme to provide free contraceptives among Mexican citizens. Between 1973 and 1979 these 'family planning' campaigns targeted potential users in urban and sub-urban zones. However, since the beginning of the 1980s 'family planning' campaigns were extended to reach rural zones (Cabrera 1994). In the last 20 years coverage increased significantly but universal access to family planning services has not yet been achieved. The program is widely considered a success in terms of the increasing diffusion and adoption of contraceptives in Mexico. In fact, while in 1976 thirty percent of all married women — or living in consensual union — were active users, in year 1998 the corresponding figure rose to almost to seventy per cent (INEGI 2001b, Gomez 1996). Since the late 1970s the public sector became the main supplier of contraceptives in the country (INEGI 2001b).

The demand for contraceptives has changed significantly in the last twenty years. At the end of the 1970s nearly 35% of all users of contraceptives in Mexico adopted the pill, 19% IUD, and 9% permanent female sterilization (PFS). In contrast, in 1998 PFS was the most popular method (51%), followed by IUD (24%) and traditional methods (10%). At this latter date, the pill was selected by less than six per cent of all active users (see Figure 1; INEGI 2001b, Gomez 1996). Various factors are behind the change in the demand for contraceptives. First, there is currently (and since late 1970s) a deliberate attempt by the public health system to promote the adoption of definitive natal control among women that have three or more children (Zavala de Cosio 1990). Second, empirical evidence suggests that Mexican women are shifting to PFS because they feel that alternative methods may either lead to serious illness — such as cancer — or result in unwanted pregnancy due to ineffectiveness (Lindstrom 1998). Finally, young individuals do not appear to be adopting contraception at the beginning of their sexual life. Instead, there is a general tendency to wait until the arrival of the first or second child to initiate the use of contraceptives (Gomez 1996).

3 Data and variable definition

Data from the National Survey of Demographic Dynamics 1997 (ENADID from its acronym in Spanish INEGI 1999) is used. The ENADID is a micro-data-set containing detailed information on Mexico's demographic dynamics that was created by the National Institute of Statistics, Geography and Informatics of Mexico (INEGI). The data set contains economic and demographic information on 88,022 Mexican women aged between 15 and 55 years. Detailed dates of birth of all women and their children were collected. Excluding observations with missing information for education and/or with unreasonable dates of birth of either the reference woman or her first child, a total of 78,467 valid cases constitute the sample for the analysis.[2]

A strict definition of a waiting time to first birth would necessarily use age at menarche as the event by which women become at risk of entering motherhood. Most fertility surveys, however, do not collect such information. That is the case of the ENADID. Following the approach of Newman and McCulloch (1984) and Heckman et al. (1985), a common event to all women is used here as opening event, or starting point, for calculating duration intervals. In particular, it is supposed that women become at risk at age 12. Previous studies have used instead age at marriage as the 'opening event'. The present paper does not follow that practise because age at marriage is potentially endogenous and its use as starting point might lead to sample selection bias (for more on this topic see Heckman 1979). The dependent variable, duration, is defined as number of years from age 12 to first birth if first birth occurred by the time of the survey, or number of years from age 12 to 1997 otherwise. A dummy variable, **firstBirth**, indicates whether a duration interval represents a completed spell (**firstBirth** = 1) or a censored observation (**firstBirth** = 0). Mean duration is around 10 years and nearly 60% of all cases report completed spells — i.e., a duration interval ended by a first birth (see Table 1). The dependent variable is constructed in terms of whole years in order to facilitate the estimation of discrete time duration models and avoid, by those means, imposing a-priori restrictions on the form of the baseline hazard.

Using information on women's date of birth three cohorts of age are defined: 1942-1962 (base group), 1963-1972 and 1973-1982. Three corresponding indicator variables were generated: **c4262**, **c6372**, **c7382**. This generation split generates three age groups that contribute approximately one third of the sample each. Notice that the base group contains cases of women born over a twenty-year period while the other two cohorts contain cases of women born over two consecutive ten-year periods. Choosing a base age group that spans a relatively long period of time is largely innocuous as the definition of 'old' and 'young' generations is somehow arbitrary. In addition, given that the Mexican government initiated its new 'family planning' campaigns in 1973, it is intuitive (and interesting) to compare groups of women that were already at risk, or about to become at risk, at the time of the policy innovation (i.e., generation 1963-1972) with groups of women that become at risk after the start of the new policy (generations 1973-82).[3]

Religion is controlled for by the inclusion of a zero/one variable splitting Catholics (**catholic** = 1) from non-Catholics (**catholic** = 0). This two-cell classification of religious groups in Mexico seems sensible given that nearly 90% of Mexicans are Catholic and a further 7% admit to be Protestant. Place of birth is taken into account by defining a set of 31 geographic dummies — 32 federal entities compose Mexico, Mexico City (D.F) is left to be the base group. To proxy ethnic group an indicator variable, **indspker**, that takes the value of one if the interviewed woman speaks an indigenous language and zero otherwise is included. Women who speak an indigenous language contribute in total a six per cent of the sample. Notice that **Indspker** proxies broad ethnic group (indigenous/mixed) rather than specific socio-cultural community. Finally, variable **Edu12** controls for education at age 12. **Edu12** is a proxy variable for skills and human capital accumulated before the onset of reproductive life. **Edu12** is bounded between zero and six and variation in the data is generated by year repetition, limited supply of education services in rural and marginal urban zones, and long-term financial difficulties of the parental household. Clearly, though children have little influence on their early education **Edu12** may be endogenous. However, as it is usual in most data sets, no valid instruments for education are available in the ENADID. Thus, **Edu12** is treated as an exogenous variable and the reader should interpret results with this in mind.

Women's income, work and marriage status are not considered here as explanatory variables. These variables are likely to be endogenous in a fertility model and their use might lead to simultaneous equation bias and invalid inference. No information on women's family background or other valid instruments for income, work and marriage status is available. Hence, estimation of a system of simultaneous equations is infeasible and the researcher should focus instead in obtaining reduced form duration models.

4 Econometric issues

Discrete time (grouped) duration models are used. Suppose initially that the *underlying* continuous-time hazard for individual i, $\theta(t)$, belongs to the Proportional Hazards family,

$$\theta(t) = \lambda(t) \exp\left[\mathbf{x_i}'\boldsymbol{\beta}\right], \tag{1}$$

where $\mathbf{x_i}$ represents a $K \times 1$ vector of characteristics for individual i (including the constant term), $\boldsymbol{\beta}$ represents a $K \times 1$ vector of coefficients to be estimated, and $\lambda(.)$ represents the baseline hazard. Notice that any negative (positive) term in the vector $\boldsymbol{\beta}$ implies a reduction (increase) of the hazard of observing a first birth (failure) and, consequently, induces a longer (shorter) mean duration. Since duration is coded in terms of years, if the vector of individual characteristics remains unchanged between time t and time $t+1$ the probability of observing a first birth at time t given that at time $t-1$ first birth had not yet occurred may be written as,

$$h_{it} = 1 - \exp\left\{-\exp\left[\mathbf{x_i}'\boldsymbol{\beta} + \gamma(t)\right]\right\} \tag{2}$$

where,

$$\gamma(t) = \ln\left[\int_{t}^{t+1} \lambda(s)ds\right].$$

Notice that no restrictions on the form of $\gamma(.)$ have been imposed.[4] This feature creates the opportunity of estimating $\gamma(t)$ using non-parametric techniques. From equation (2) the reader may conclude that the discrete time hazard, h_{it}, takes an Extreme Value form.[5] More importantly, Model (2) can be thought of as a sequence of non-identical Bernoulli trials where each individual contributes one observation per survived period (for further details see Narendranathan and Stewart 1993). Using this fact the overall contribution of the $i-th$ individual to the likelihood can be written as,

$$L_i = \prod_{t=1}^{d_i}\left\{\left[h_{it}\right]^{w_{it}}\left[1-h_{it}\right]^{1-w_{it}}\right\}, \tag{3}$$

where w_{it} is a binary variable taking one if the duration spell for the $i-th$ individual is a non censored observation and duration, d_i, equals t. Interpreting the discrete time duration model as a sequence of binary choice regressions creates the opportunity of relaxing the proportional hazards assumption and adopting alternative binary choice models for h_{it} such as Logit and Probit. Logistic and Normal hazards exhibit a symmetric distribution and constitute a valuable contrast to the skewed Extreme Value distribution.[6] This sort of model has been estimated in the field of labour economics by various authors including Meyer (1990), Sueyoshi (1995), and Arulampalam and Stewart (1995). In the present work Probit hazards were estimated along with the benchmark EV hazards in order to assess the robustness of results to the proportional hazards assumption.

To accommodate unobserved heterogeneity a random term, v_i, is introduced into equation (1). This random term captures differences in the hazard induced by heterogeneity in omitted characteristics such as fecundity and/or skills. Unobserved heterogeneity is supposed to be orthogonal to the vector of explanatory variables $\mathbf{x_i}$. Identification is secured if at least one continuous variable is included into the design matrix $\mathbf{x_i}$ (for further reference see Elbers and Ridder 1982). Conditional on unobservables the hazard in equation (2) becomes,

$$h_{it}(v_i) = 1 - \exp\left\{-\exp\left[\mathbf{x_i}'\boldsymbol{\beta} + \gamma(t) + v_i\right]\right\}. \tag{4}$$

On the basis of equation (4) the likelihood function might be re-written in terms of the conditional hazard $h_{it}(v_i)$ and the probability density function of v_i, $f(v_i)$,

$$L_i = \int \left\{ \prod_{t=1}^{d_i} \left\{ \left[h_{it}\right]^{w_{it}} \left[1 - h_{it}\right]^{1-w_{it}} \right\} \right\} f(v_i)\, dv_i. \tag{5}$$

The model is closed once a functional form for the mixing density $f(v_i)$ is chosen.[7] Normal and Gamma densities are a common choice. Alternatively, as suggested by Heckman and Singer (1984), a non-parametric maximum likelihood estimator (NPMLE) can be used to estimate an empirical approximation to $f(v_i)$. This method is based on results establishing that, given a functional form of the hazard, a NPMLE can be approximated by a finite discrete mixture — or mass points. This methodology is used in the present work.

A further issue to be addressed is that some individuals may be sterile or dislike children to such an extent that they end their fertile life childless. Those individuals are never at risk of entering parenthood and, consequently, report extremely long duration spells. In order to account explicitly for lifetime childlessness the model is extended by means of introducing a mass point at minus infinity (referred simply as 'final mass point' in future discussion). This feature allows for a 'mover-stayer' framework. The $i-th$

individual contribution to the likelihood function is therefore:

$$L_i^* = \frac{\delta}{1+\delta}\left\{\prod_{t=1}^{d_i}(1-w_{it})\right\} + \frac{L_i}{1+\delta}, \tag{6}$$

with L_i as in equation (5). The basic idea behind model (6) is that a large negative value of the unobservable v_i induces an extremely long duration interval and, consequently, increases the likelihood of observing a censored observation. This model is equivalent to the Split-Population model of Schmidt and Witte (1989).

Given that the alternative specifications considered in the present study are non-nested, the Akaike information criterion (AIC) and its consistent version (CAIC) are used to compare them. Model selection is also performed on the basis of a Lamer-Schwarz metric (LS). If AIC is used for selection, a best fitting model will be the one that achieves the minimum AIC against all other considered alternatives. CAIC and LS work in a similar fashion. All these information criteria penalise for the loss of degrees of freedom when an additional parameter is included and are widely used for model selection in many fields of applied work. Formally the statistics are defined as:

$$\begin{aligned} AIC &= -2ln(L) + 2k; \\ CIAC &= -2ln(L) + \{ln(n) + 1\}; \\ LS &= ln(L) - \left(\frac{ln(n)}{2}\right)k. \end{aligned}$$

5 Empirical results

Table 3 presents empirical results. Column 1 reports estimates for a EV hazard while column 2 reports estimates for a Probit hazard. As discussed earlier, the latter specification removes the proportional hazards assumption and constitutes a valuable contrast to the benchmark EV hazard. In both specifications unobserved individual heterogeneity is controlled for by the inclusion, and empirical estimation, of two mass points for approximating the distribution of unobservables. Finally, a final mass point explicitly accounts for the presence of a lifetime childless group of individuals. Including an additional mass points did not produce significant improvements in the log-likelihood and the vector of coefficients on explanatory variables remained unchanged. During estimation other specifications for unobserved heterogeneity were intended. In particular, EV and Probit hazards with no unobserved heterogeneity and with unobserved heterogeneity but no final mass point were estimated. Similarly, EV and Probit hazards with Normal unobserved heterogeneity and final mass point were also obtained. Table 3 contains the best fitting models for each type of hazard function considered here.

Notice first that except for the probability mass attached to the final mass, coefficients in columns 1 and 2 are not directly comparable — comparison should rather be done in terms of the implied average partial effects (APE) reported in Table 4. It is remarkable though that coefficients on explanatory variables have the same sign across the two versions of the hazard. Moreover, in both specifications coefficients are different from zero at all conventional significance levels. In a similar fashion, the estimated location for the first mass, $mass1$, and the probability attached to it, $Pr(mass1)$, are both highly significant. In other words, unobserved heterogeneity is effectively present.

A Probit hazard seems to fit better the data. In fact, comparing columns 1 and 2 results in Table 3 the reader may find that a Probit hazard attains a sensibly lower log-likelihood than the EV hazard specification (-158,294.17 and -158,208.50 for EV and Probit respectively). Model selection on the basis of AIC, CAIC and LS statistics suggest as well that a Probit hazard outperforms the Extreme Value specification. Relaxing the proportional hazard assumption seems therefore the best strategy in the present context. In either case, EV or Probit, the probability attached to the final mass is around 0.12. This finding suggests then that 12% of Mexican women are expected to remain childless for their entire lifetime.

Figure 2 presents estimates for the average hazard based on results from Table 3. Like in the case of APE, for any potential realisation of the heterogeneity term v the conditional hazard of failure at each duration time, $h_t(v)$, was calculated. Next, the average hazard was obtained taking the expected value of $h_t(v)$ over the distribution of the random term v, $f(v)$. This procedure delivers a series of points — one for each discrete duration interval where at least a single failure was reported — that are then plotted. Graphs in Figure 2 represent therefore a non-parametric estimator of the average hazard. All calculations were performed for a typical individual.[8]. Duration dependence exhibits the same pattern for either EV or Probit. First, at the beginning of the duration spell, the hazard increases with duration time. Then, with the passing of time, the hazard becomes flat before it starts to exhibit negative duration dependence. This functional form for the hazard conforms the intuition that the cost of first pregnancy is high at the beginning of fertile life, decreases with the passing of time, and then, at a certain point, starts to rise again as the end of fertile life is approached.

From Figure 2 the reader may conclude that the average hazards from the EV and Probit models look fairly similar for short intervals. For long intervals, however, important differences are detected. While the EV hazard peaks around duration 20, a Probit hazard peaks three years later — at duration 23. Moreover, the Probit hazard for durations longer than 20 is found to be systematically higher than the EV hazard. In other words, the Probit hazard predicts a larger proportion of failures at the end of fertile life than an EV hazard would predict. This feature gives an intuition of why a Probit hazard was found to fit the data better than an EV hazard. Both EV and Probit hazard detect an acute peak in the hazard function at duration 27; that is, at age 39. This final peak is probably associated with reductions in women's contraception effort at the very end of fertile life, maybe because the risk of a pregnancy is wrongly under-estimated with the detection of temporary losses of fecundity. Alternatively, the final peak in the hazard might indicate the existence of a group of women who wait until the last periods of their fertile life to make decisive efforts to enter motherhood, just as predicted by Happel et al. (1984).

Concentrating attention in the best fitting model (i.e., Probit hazard) results suggest that Catholic individuals have a significantly lower risk of first birth than non-Catholic individuals (and consequently a longer mean duration). In fact, the probability of failure at mean duration for a Catholic individual is estimated to be on average 1.57% lower that the corresponding probability for a non-Catholic individual. In other words, the average partial effect (APE) on **Catholic** at mean duration is approximately -0.0157 (see column 2 of Table 4). At first sight this result might seem counter-intuitive given that the Catholic Church traditionally opposes the adoption of contraceptives as a way to regulate fertility, an attitude that is generally thought to increase the odds of unwanted pregnancy among catholic individuals. However, the result is better understood if the reader considers that, besides being reluctant to adopt modern birth control, the Catholic Church also discourages sexual activity out-of-wedlock, an attitude that in many cases leads to the delay of first sexual intercourse. Hence, the net effect of Catholicism on the hazard may well be negative rather than positive. In the case of Mexico, where Catholic individuals represent nearly 90% of the population and a further 9% are either Protestant or Atheist (INEGI 1999), it is intuitive that non-Catholic individuals may initiate sexual activity earlier than Catholic individuals as Protestants and Atheists are traditionally more liberal about such issues. Therefore, though Catholic youngsters may be highly reluctant to adopt modern contraception, they still may have lower sexual activity at early ages than their non-Catholic peers.[9]

Being an Indian language speaker, in contrast, implies an increased hazard of failure of almost three percent, for the average partial effect on **indspeker** at mean duration is 0.0288. This effect is the synthesis of a number of factors. Most Mexican Indians live in rural zones on settlements of less than 1,500 inhabitants that are far from the main regional cities and where the main economic activity is agriculture. The majority of indigenous individuals have a small extension of land and produce mainly for self-consumption. As a general rule, these indigenous zones have very limited supply of education and health services. Hence, commonly, Indian individuals do not study beyond the basic instruction available in their localities — which in many cases is below six grades. Post basic education is even more limited in rural and indigenous zones despite the fact that from 1996 it became compulsory for all Mexican children. This is one of the main factors that keep average education among indigenous individuals at 5.38 years, well below the 8.13 years corresponding to Mexico as a whole. Limited health services are associated with a relative high rate of infant mortality among Indians, which in 2000 reached 38.5 deaths per 1,000 births, above the national rate of 17 deaths per 1,000 births (CONAPO 2002). Culture might also play some role in the relatively high hazard of failure among Mexican Indians, especially if tradition keep women outside the labour market (that is, far from paid jobs) and reduce their bargaining power inside the household. From the present analysis, however, there is no way of inferring the relative importance of these cultural factors as determinants of early motherhood.

As economic intuition would suggest, education at age 12 is found to reduce the hazard of a first pregnancy (or equivalently, to increase mean duration). In fact, the average partial effect at mean duration on **Edu12** a value of -0.0245. In other words, an extra year of schooling at age 12 reduces the likelihood of first birth by approximately 2.5 points. These results are consistent with previous findings reported by Newman and McCulloch (1984) for the case of Costa Rica.

Regarding generational effects Table 3 shows that women in cohort 1963-1972 have a higher risk of first birth than women in the base age group (1942-1962). The implied average partial effect at mean duration is though just about 0.0076, indicating a slight reduction on mean duration. This result implies that women born between year 1963 and year 1972 accelerated their entrance to motherhood in relation to women in the base group. The effect, however, is rather small. In the case of the youngest group results strongly support the hypothesis that women born between 1963 and 1982 delayed first birth relative to the experience of women in the base age group. This finding is clearly witnessed by the negative and significant coefficient on **c7382** in column 2 of Table 3. The implied average partial effect at mean duration attains a value of -0.0131, indicating a reduction on the risk of first birth of around 1.3 points (and thus and increase on mean duration) in reference to the 1942-1962 age group. Though the delay effect is rather small, it is highly significant.

Up to now the analysis has maintained the assumption that entry at different calendar times lead to parallel shifts in the hazard function. Further, explanatory variables have been constrained to affect the hazard exactly in the same fashion across different age groups. Finally, a unique distribution for the unobservables has been estimated for individuals belonging to different age cohorts. These assumptions are stringent and unlikely to be met in practice. Hence, important insights could be gained by estimating a separate hazard function for each cohort.

Following a strategy similar to that previously used, various forms of the hazard were specified for each cohort. In all cases model selection on the basis of AIC, CAIC and LS indicate that a Probit hazard is the best fitting model. Regarding unobserved heterogeneity, selection on the basis of AIC, CAIC and LS suggest a model with two plus final masses for cohorts 1942-62 and 1963-1972, while a model with exclusively the final mass was supported for cohort 1973-1982. Table 5 reports APEs calculated on the basis of the best fitting hazard model for each age group. It appears first that, though limited, variation in the APE across the different cohorts is non-negligible given the size of the APEs obtained. **Catholic**, for instance, is associated to a 1.63 point reduction in the likelihood of failure for women born between 1942 and 1962. In comparison, **catholic** in the 1963-1972 group leads to a decreased risk of first birth of around 1.80 points. This implies that the Catholic/Non-Catholic relative risk of failure in the later group is 1.43 lower than in the former group. If cohorts 1942-62 and 1973-82 are contrasted instead it is the younger and not the older generation who bear an increased Catholic/Non-Catholic relative risk of first birth. A similar story describes variation of the APEs of **indspker** and **Edu12** across age groups. In the case of **indspker**, however, there is a clear pattern indicating that the passing of calendar time has led to increments in the risk of first birth of indigenous language speakers in relation to non-indigenous language speakers.

Estimated average hazards for each cohort are reported in Figure 3. As in the case of APEs, calculations are based on the best fitting model for each age-specific group. Various pieces of new information are obtained. First, there is clear empirical evidence that younger cohorts of women are delaying first birth relative to older cohorts. This is witnessed by an almost everywhere inwards shift of the **c6372** hazard line with respect to **C4262** line, and by the inwards shift of the **c7382** with respect to the **c6372** line. Second, inspection of Figure 3 shows that the assumption that entry at different calendar times leads exclusively to parallel shifts in the hazard is hardly supported by the data. This implies that misspecification of the hazard may have misled inference in previous paragraphs. In fact, in the light of the new evidence, it is possible to conclude that women in both cohorts 1963-1972 and 1973-1982 delayed significantly first birth with respect to the experience of women in the base age group.

6 Summary and conclusions

In the last twenty years female permanent sterilisation (FPS) has become the most popular contraceptive method in Mexico. During the same period of time the demand for pills, condoms and other short-term contraceptives dropped sharply and consistently. The changes in the demand for contraceptives have raised concerns among demographers that the timing of children in Mexico might remain unchanged despite the reduction of fertility rates in the last few decades. In particular, it is thought that women might be cutting off fertility at high parities without modifying their reproductive calendar. If present, this behaviour maintains a traditional patten of early entry into motherhood and limits the opportunities of Mexican women to accumulate human capital and improve their standards of life. Moreover, by deterring accumulation of human capital, sustained early entry to motherhood may also limit the economic growth of Mexico as a whole.

The present paper shows that despite the popularity of PFS and the drop in the demand for short-term contraceptives, young generations of Mexican women have effectively delayed first birth. Hence, the hypothesis that Mexican women are following a fertility cutting off strategy that does not modify their reproductive calendar is not supported by the empirical evidence.

The study finds that Catholic individuals have lower hazard of entering motherhood than non-Catholic individuals. A result that supports the idea that Catholicism has not been a relevant factor preventing women from modifying their fertility behaviour in Mexico. Instead, Catholicism seems to play a relevant role for the delay of first birth. From the analysis is not possible to infer the reasons behind this result. But a postponement of marriage and sexual abstinence previous to it seems to be a plausible explanation. Finally it is found that indigenous language speakers enter motherhood faster than non-indigenous language speakers, a result that is intuitive given that, as a general rule, individuals belonging to the indigenous ethnic groups of Mexico have more limited access to education and health.

Appendix

Table 1. Descriptive statistics

Variable	Description	Mean	Std. Dev.	Min	Max
Age	age in years.	29.66	10.47	15	55
duration	See note.	9.94	6.41	3	43
firstbirth	=1 if first birth observed; 0 otherwise.	0.60	-	-	-
Catholic	=1 if Catholic; 0 otherwise.	0.89	-	-	-
indspker	=1 if indian language speaker; 0 otherwise.	0.06	-	-	-
Edu12	education at age 12 in years.	5.36	1.38	0	6
c4262	Cohort base group				
c6372	=1 if born within 1963-1972; 0 otherwise.	0.30	-	-	-
c7382	=1 if born within 1973-1982; 0 otherwise.	0.38	-	-	-
+ 32 birthplace dummies (base Mex. City).					
Number of observations					78,467

Note: Duration is defined as years between age 12 and firth birth if a completed spell is observed or years between age 12 and 1997 otherwise.

Figure 1. Demand for contraceptives in Mexico

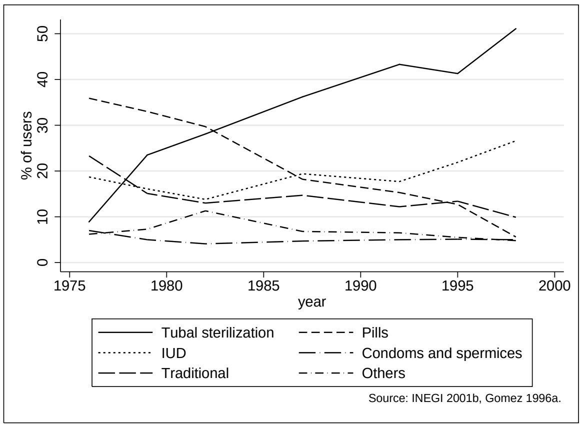

Table 2. Descriptive statistics–Cohorts

Variable	Obs	Mean	Std. Dev.	Min	Max
Cohort 1942-1962 (base)					
Age	25,454	42.42	5.54	34	54
duration	25,454	12.92	8.78	3	43
firstbirth	25,454	0.85	0.35	-	-
Catholic	25,454	0.90	-	-	-
indspker	25,454	0.06	-	-	-
Edu12	25,454	4.84	1.75	-	-
Cohort 1963-1972					
Age	23,214	29.06	2.92	24	35
duration	23,214	10.86	4.64	3	22
firstbirth	23,214	0.75	0.43	-	-
Catholic	23,214	0.89	-	-	-
indspker	23,214	0.06	-	-	-
Edu12	23,214	5.50	1.23	-	-
Cohort 1973-1982					
Age	29,799	19.24	2.83	15	25
duration	29,799	6.67	2.56	3	12
firstbirth	29,799	0.26	-	-	-
Catholic	29,799	0.89	-	-	-
indspker	29,799	0.06	-	-	-
Edu12	29,799	5.71	0.95	-	-

Figure 2. Estimated hazard

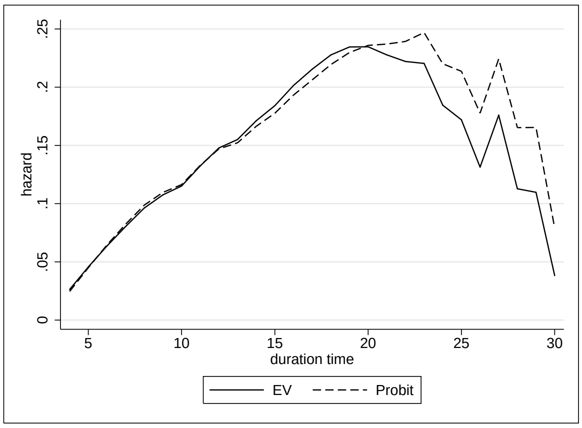

Table 3. Empirical Results — Hazard function

	(1) Extreme Value	(2) Probit
Constant	-3.8830 [0.0500]**	-2.9860 [0.0547]**
Catholic	-0.1484 [0.0201]**	-0.0802 [0.0109]**
indspker	0.2625 [0.0281]**	0.1471 [0.0153]**
Edu12	-0.2203 [0.0050]**	-0.1233 [0.0028]**
c6372	0.0732 [0.0146]**	0.0387 [0.0079]**
c7382	-0.1398 [0.0168]**	-0.0668 [0.0088]**
Birthplace dummies	Yes	Yes
mass1	-1.6649 [0.0575]**	1.0744 [0.0464]**
Pr(mass1)	0.1708 [0.0130]**	0.7706 [0.0105]**
Pr(massend)	0.1220 [0.0040]**	0.1256 [0.0042]**
Log-likelihood	-158,294.17	-158,208.50
χ^2	16,209.13	17,008.8
Pr $> \chi^2$	0.0000	0.0000
AIC	316,724.34	316,547.0
CAIC	317,422.73	317,214.58
LS	-158,677.36	-158,574.79

Note: NPMLE estimates (Two plus final mass points unobserved heterogeneity). Standard errors in brackets, **(*) indicates significance at 1% (5%).

Figure 3. Estimated hazard–Cohort of birth.

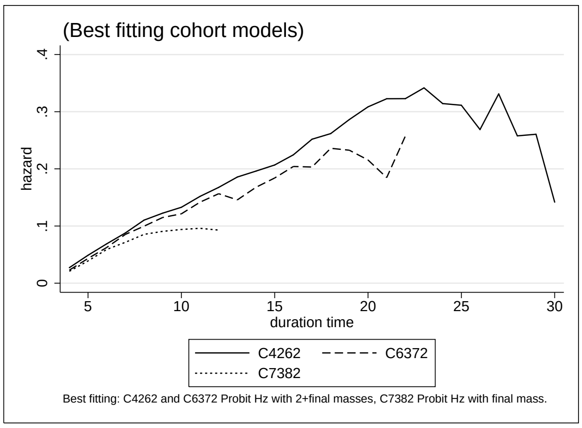

Table 4. Average partial effects

	Extreme Value	Probit
Catholic	-0.0167**	-0.0157**
indspker	0.0350**	0.0288**
Edu12	-0.0264**	-0.0245**
c6372	0.0090**	0.0076**
c7382	-0.0158**	-0.0131**

**(*) indicates significance at 1% (5%) of the estimated coefficient.

[1] If $d_j(\mathbf{x_i}, v_i)$ represents the partial effect of x_j on the conditional hazard, given the vector of observed variables $\mathbf{x_i}$ and the unobserved random effect v_i, the average partial effect of x_j is calculated as the expected value of $d_j(\mathbf{x_i}, v_i)$ over the density function of the random effect $f(v_i)$.

[2] Average Partial effects of continuous variables are calculated at the mean value, mean duration,and setting all dummies to zero. Average Partial effects for discrete variables are calculated by the difference measure at the mean of all continuous variables, mean duration, and setting all other dummy variables to zero.

Table 5. Average partial effects — Cohort of Birth

	1942-1962	1963-1972	1973-1982
Catholic	-0.0163**	-0.0180**	-0.0102**
indspker	0.0161**	0.0291**	0.0370**
Edu12	-0.0244**	-0.0308**	-0.0243**

Based on the best fitting model for each cohort. **(*) indicates significance at 1% (5%) of the estimated coefficient.

[1] If $d_j(\mathbf{x_i}, v_i)$ represents the partial effect of x_j on the conditional hazard, given the vector of observed variables $\mathbf{x_i}$ and the unobserved random effect v_i, the average partial effect of x_j is calculated as the expected value of $d_j(\mathbf{x_i}, v_i)$ over the density function of the random effect $f(v_i)$.

[2] Average Partial effects of continuous variables are calculated at the mean value, mean duration,and setting all dummies to zero. Average Partial effects for discrete variables are calculated by the difference measure at the mean of all continuous variables, mean duration, and setting all other dummy variables to zero.

Endnotes

1. The total fertility rate (TFR) is a measure of the number of children that a woman would have at the end of their fertile life if she follows the current 'typical' fertility behaviour at all stages of her life (it can be expressed in terms of children per 1,000 women). Age-specific fertility rates indicate the number of birth per 1,000 women in different age-specific groups.
2. Cases where the given dates of birth of mothers and their children implied a negative duration interval were excluded. The analysis is done conditional on this selection.
3. During estimation it was found that moving the calendar limits of the generation dummies for +5 and -5 years did not result in important changes in the estimated parameters and their standard errors.
4. The *underlying* continuous time hazard $\theta(.)$ is the model the researcher would estimate if continuous-time duration data were available. However, if one collects duration intervals that only change in a discrete manner (say, years) and $\mathbf{x_i}$ is constant between t and $t+1$, then $\theta(.)$ can be written as h_{it}. In the duration data literature h_{it} is known as the *discrete* time hazard function (see for instance Meyer 1990, Sueyoshi 1995).
5. Notice that the Extreme Value distribution of the discrete time hazard h_{it} is a direct consequence of the proportional hazards functional form of the underlying continuous time hazard $\theta(t)$. To avoid confusion between h_{it} and $\theta(t)$, h_{it} is referred to as the 'EV hazard' whenever $\theta(t)$ is supposed to belong with the proportional hazards family.
6. The Extreme Value type I distribution may not be attractive in applied work because it has a fat right tail (i.e., the distribution is skewed to the right and has $skewness = 1.13955$). Clearly, for large samples central limit theorem arguments would suggest that a rather symmetric distribution such as Normal is appropriate.
7. To simplify exposition, from now on, the reader should understand 'discrete time' hazard anytime the hazard function is referred to.
8. The typical individual was found to be Catholic and had 5.35 years of education at age 12.
9. According to the ENADID, between 1992 and 1997 single mothers contributed 5% of the most recent pregnancies that resulted in live births. No single mothers are reported among non-Catholic women who were born between 1953 and 1957. In contrast, the generation of non-Catholics who were born between 1968 and 1972 contributed 8.41% of children born to single mothers. In other words, non-Catholic women in the young cohort contributed a higher proportion of out of wedlock births than non-Catholic women in the older cohort. To complete the picture data from the ENADID shows that while 78% of women in the 1953-1957 cohort married

before age 25 (or entered a consensual union), only 69% of women in the 1968-1972 cohort married before that age. Clearly, these descriptive statistics support the hypothesis that, at least in relative terms, the Catholic Church is succeeding in persuading young couples to delay marriage and sexual intercourse.

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