Child Mortality Estimation by Time Since First Birth

Kenneth Hill
Department of Population and Family Health Sciences
Johns Hopkins University

and

Maria-Elena Figueroa
Center for Communications Programs
Johns Hopkins University

Introduction

The best-known and most widely-applied estimation method developed by William Brass is the measurement of child mortality from the proportions dead of children ever borne by women classified by age group (Brass;1964,1975). The basic principle of the method is that age of the mother can serve as a proxy for the exposure time of her children, so that the proportion dead for women of a given age group can be converted into a defined probability of dying for their children. This method had been applied to census and survey data from all parts of the developing world, and has been found to work remarkably well in a wide variety of settings. The original method has been extended by a number of authors, notably Sullivan (1972), who extended the method to groups of women classified by duration of marriage; Feeney (1975), who showed that under conditions of steadily changing mortality the estimates of mortality for particular age groups of women could be related to specified time points prior to the survey; and Trussell (1975;United Nations,1983), who expanded the model base of the estimation methods. As used today, the proportion dead of children ever borne for each group of women provides an estimate of child mortality, usually converted into the probability of dying by age 5, and a time reference for that estimate in years before the survey.

One of the key assumptions underlying the original Brass method and its extensions is that child mortality risks are uniform across the classificatory variable (age or duration of marriage of the mother) being used to proxy exposure to risk of the children. Widespread application of the age of mother based method has shown that this assumption is not valid. The child mortality estimates based on reports of younger women, particularly those aged 15-19 but also often those aged 20-24, are almost always higher than the estimates based on reports of older women. This pattern results from a real age effect, whereby children of young mothers have elevated mortality risks, and also from a selection effect, whereby women of lower socio-economic class tend to start childbearing early, and have children exposed to above average mortality risk. It is unfortunate that it is the two youngest age groups of mothers that are most affected by this bias, since it is these two age groups that reflect the most recent child mortality experience, and provide estimates with reference dates closest to the time of the survey. The youngest age groups of mothers also have the smallest numbers of children ever borne, so random errors are largest for the estimates based on them. The duration of marriage based method is less affected both by the age at childbearing selection bias and by small numbers of events, since childbearing typically occurs rapidly in the first five years of marriage across all social classes. The duration of marriage based method may, however, be affected by another selection bias, by marital status, in countries where substantial proportions of children are born outside formal unions. The experience of these children would not be captured in the duration of marriage based method until the mother married, and would then be captured at an inappropriate value of the proxy exposure indicator.

This paper develops and illustrates the application of a third approach, which uses time since first birth as the proxy for exposure. This approach will be little affected by socio-economic selection bias, but will be applicable to populations in which substantial proportions of children are born outside formal unions. The method is developed by using fertility and mortality models to simulate children ever borne and children dead for women in five year time since first birth groups, and then to relate the proportions dead to standard mortality measures allowing for different fertility patterns. For the method to be applied, additional questions on month and year of first birth must be included in the census or survey. The number of surveys that have included such questions without collecting a full birth history is small, so the number of applications currently available is also small.

Fertility by Time Since First Birth

Fertility by time since first birth is modelled in a manner analogous to that shown by Rodriguez and Cleland (1988) to fit patterns of fertility by time since first union closely. The model assumes a time dependent pattern of natural fertility at time since first birth t, n(t), which is modified by two parameters, one determining the level of fertility, the other determining the extent of deviation over time of the actual pattern from the natural pattern. The model is

$$f(t) = n(t) \exp(a + \beta t) \qquad (1)$$

where a is the level parameter and ß is the time-dependent deviation.

The natural fertility pattern was derived from data for the Hutterites. Births and exposure by single year time since first birth categories were calculated, and single year rates computed. Table 1 shows the reported rates by single year, and the rates after applying a LOWESS smoothing procedure to the values for periods since first birth greater than three years (no smoothing is applied to the rates for the first three years, since they fluctuate sharply reflecting the effects of nonsusceptibility).

The model described in equation (1) and the Hutterite standard was then applied to data from 31 Demographic and Health Surveys. Fits by and large were satisfactory, with values of a ranging from -1.05 to -1.97, and values of ß from -0.008 to 0.082. Figure 1 shows plots of $log(f(t)/n(t))$ against t for two countries, one with a low value of ß (Thailand) and one with a high value

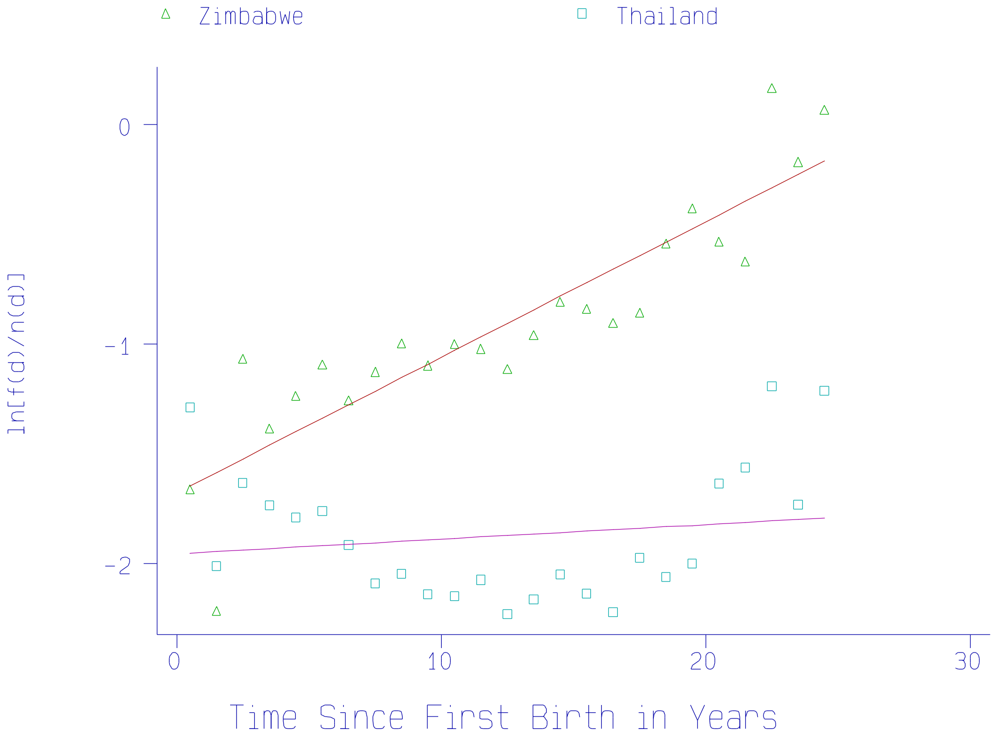
TABLE 1: Observed Hutterite and Smoothed Fertility Rates by Year Since First Birth

Time Since First Birth (a,a+1)	Hutterite Birth Rate	Smoothed Birth Rate
0	.0450	.0450*
1	.8432	.8432*
2	.5592	.5592*
3	.6007	.5947
4	.5783	.5742
5	.5285	.5545
6	.5488	.5365
7	.5269	.5201
8	.4813	.5039
9	.4767	.4870
10	.4720	.4695
11	.4577	.4517
12	.4825	.4329
13	.4201	.4126
14	.3692	.3903
15	.3807	.3653
16	.3476	.3371
17	.3592	.3065
18	.2553	.2744
19	.2580	.2414
20	.2054	.2074
21	.1726	.1717
22	.1025	.1349
23	.1239	.0986
24	.0623	.0644

* Values for 0,1 and 2 years are not smoothed.

of ß (Zimbabwe). Examination of the 31 applications showed no systematic relation of a to the overall level of fertility, but did suggest that ß tended to be small in low fertility countries and high in higher fertility countries. This relationship was taken into account in choosing combinations of fertility level (total fertility rate) and pattern (a and ß) in the simulations of proportions dead of children ever borne. It is an important characteristic of this model that the first birth is automatically excluded. In the simulations, the first birth is exposed to mortality risk for the maximum exposure time possible, while subsequent births are distributed by exposure according to the model. For this reason, fertility level, determining the ratio of first to subsequent births, matters in these simulations, whereas it did not matter in either the age or time since first marriage based versions of the Brass method.

FIGURE 1. Fits of Fertility Since First Birth Model to Thailand and Zimbabwe



Values of a used in the simulations were -1.0,-1.5 and -2.0. As noted, values of ß varied with the level of overall fertility. On the assumption that the methodology would be most widely applicable to high fertility settings, the levels of total fertility chosen were 4.0, 5.0 and 6.0. The values of ß were 0.00, 0.02 and 0.04 for total fertility equal to four, 0.02, 0.04 and 0.06 for total fertility equal to five, and 0.04, 0.06 and 0.08 for total fertility equal to six.

Mortality Risks

Mortality risks by single years of age were taken from Coale-Demeny model life tables (Coale and Demeny;1983) for three mortality levels (10, 15 and 20, corresponding to expectations of life at birth for females of 42.5, 55 and 67.5 years respectively) for each of the four families of tables. Single year values of the survivors to exact age x, l(x), for values of x from 5 to 25 were obtained by linear interpolation, and a both-sex life table was obtained by combining the male and female tables using a sex ratio at birth of 105 males per 100 females. Survivorship probabilities for single years of age were obtained by assuming linearity of the survivorship curve across one year periods above age one, and by applying an average age at death of 0.3 years for children dying under one for levels 10 and 15, and of 0.2 years for level 20.

Population Growth Rate

The fertility model used in the simulations varies with the total fertility rate, so the population growth rate, used to weight simulations at different times since first birth into five year groups, also varies with the fertility pattern chosen. The growth rate for a particular simulation was estimated assuming a mean length of generation of 29 years, and an estimate of the Net Reproduction Rate derived from the TFR as (TFR/2.05)(l(29)/l(0)), where l(29) and l(0) are life table survivors to age 29 and 0 respectively.

Simulating Proportions Dead Among Children Ever Born

The number of children ever borne and the number dead among those children is estimated at single year points since the first birth. The number of children ever borne at time since first birth *a*, CEB (a), is simply one (the first birth) plus the sum of the time since first birth fertility rates up to time *a*:

$$\mathrm{CEB}(a) = 1 + \sum_{0}^{a-1} \mathrm{f}(t) \qquad (2)$$

where f(t) is the fertility rate at t years after first birth. The number of dead children at time since first birth *a*, cd (a), is the sum of the products of fertility rates and probabilities of dying:

$$\mathrm{CD}(a) = \mathrm{q}(a) + \sum_{0}^{a-1} \mathrm{f}(t)\,(1 - {}_1\mathrm{L}_{a-t}/\mathrm{l}(0)) \qquad (3)$$

where q(a) is the probability of dying by age a, ${}_1L_{a-t}$ is the life table person-years lived at age (a-t), and l(o) is the radix of the life table. The population growth rate, r, is then used to combine average numbers of children ever borne and children dead for five year time since first birth intervals, on the assumption that the number of women at each time since first birth category is given by exp(-r**a*) relative to the number having a first birth. The proportion dead of children ever borne is then obtained as the ratio of CD(i)/CEB(i) for time since first birth group i. The average parity for the group is simply the average CEB(i).

Relating Proportions Dead to Standard Mortality Indices

Proportions dead in a particular interval are affected by mortality risk and by the time distribution of births. A proportion dead is converted into an estimate of a probability of dying by an exact age of childhood by allowing for the time distribution of births. Ratios of average parity are used in this conversion, such that

$${}_nq_0/\mathrm{PD}(i) = a(i) + b(i)\{P(1)/P(2)\} + c(i)\{P(2)/P(3)\} \qquad (4)$$

where P(1), P(2) and P(3) are the average parities in each five-year time since first birth interval. Linear regression was used to estimate values of a(i), b(i) and c(i) from the simulated proportions dead and average parities in time since first birth groups from 0 to 4 through 20 to 24. Separate regressions were run for each family of Coale-Demeny model life table. Values of a(i), b(i) and c(i) are shown in Table 2.

The form of equation (4) is a corollary of the equations used for the age- and duration of marriage-based methods developed by Trussell (1975). Some experimentation with alternative forms was tried before adopting equation (4), however. First was the question of what indicator of fertility timing to use. The method uses time since first birth as the surrogate for child exposure time, but the first birth is exposed throughout the period, and only subsequent births are distributed according to the model. The model pattern might, therefore, be reflected better by ratios of {parity - 1} than by parity. This formulation was tried, but resulted in substantially less good fits for equation (4) than using parity.

A second question is what values of n to use in the ${}_nq_0$. Again, experimentation showed that the use of values of n of 2, 3, 5, 10 and 15 for the time since first birth groups 0 to 4, 5 to 9, 10 to 14, 15 to 19 and 20 to 24 respectively gave satisfactory results. Average exposure time of children for groups of women defined by time since first birth is somewhat longer than average exposure time for groups defined by duration of marriage, but the same relations work well for both.

Reference Dates of the Estimates

Under conditions of changing mortality, each proportion dead identifies an ${}_nq_0$ from a period life table for some point in the past, where the point in the past lies between the time of the survey and the longest exposure time of the children. Each period life table in the past is effectively weighted by the number of child deaths that occur in the period, so the reference period of an estimate can be approximated by the average time ago of the deaths of the children reported on. The reference period will be affected to a minor extent by the direction and pace of mortality change. A rapid increase in mortality will bring the reference period closer to the present, whereas a rapid decline will move the reference date further into the past. However, as Coale and Trussell (1977) show, the effect of such change is small relative to the impact of different time distributions of exposure.

The time reference of the estimates based on proportions dead classified by time since first birth is estimated from equation (5):

$$t^*(i) = d(i) + e(i)\{P(1)/P(2)\} + f(i)\{P(2)/P(3)\} \qquad (5)$$

where $t^*(i)$ is the number of years before the survey to which the estimate applies. Values of d(i), e(i) and f(i) are given in Table 3 for each family of Coale-Demeny model life tables. These reference points tend to be longer in the past for the time since first birth method than for the duration of marriage approach because the first birth is exposed to the maximum exposure period.

Table 2 Coefficients for Application of Time Since First Birth Mortality Estimation

$$\frac{{}_{n}q_{o}}{PD(i)} = a(i) + b(i)\frac{P(1)}{P(2)} + c(i)\frac{P(2)}{P(3)}$$

		Time Since First Birth (Years)				
		0-4	5-9	10-14	15-19	20-24
	n	2	3	5	10	15
North Model						
	a	1.1809	1.1298	1.2037	1.2933	1.3240
	b	-0.0787	-0.1609	-0.0107	0.0987	0.1557
	c	-0.0182	-0.0746	-0.2856	-0.4629	-0.5678
South Model						
	a	1.1697	1.1337	1.2977	1.4860	1.5278
	b	-0.1431	-0.2625	-0.0416	0.1566	0.2619
	c	0.0034	-0.0859	-0.4335	-0.7602	-0.9135
East Model						
	a	1.2023	1.1669	1.2501	1.3243	1.3279
	b	-0.1322	-0.1911	-0.0003	0.1220	0.1701
	c	-0.0022	-0.0943	-0.3383	-0.5121	-0.5770
West Model						
	a	1.1882	1.1410	1.2417	1.3631	1.4240
	b	-0.1063	-0.1953	-0.0231	0.1104	0.1934
	c	-0.0098	-0.0822	-0.3390	-0.5766	-0.7420

Table 3 Coefficients for Estimating the Time Reference of Estimates

$$t^{*}(i) = d(i) + e(i)\frac{P(1)}{P(2)} + f(i)\frac{P(2)}{P(3)}$$

Time Since First Birth (Years)

$$t^*(i) = d(i) + e(i)\,\frac{P(1)}{P(2)} + f(i)\,\frac{P(2)}{P(3)}$$

	0-4	5-9	10-14	15-19	20-24
North Model					
a	1.71	2.16	0.66	-1.96	-3.85
b	1.07	4.36	3.50	-0.90	-6.42
c	-0.35	0.12	6.65	17.66	28.94
South Model					
a	1.68	2.29	1.19	-1.01	-2.68
b	0.96	3.84	3.45	-0.18	-5.06
c	-0.32	-0.01	5.41	15.03	25.21
East Model					
a	1.68	2.19	0.71	-1.96	-4.06
b	0.99	4.28	3.63	-0.71	-6.35
c	-0.33	0.02	6.36	17.42	29.14
West Model					
a	1.70	2.20	0.86	-1.46	-2.97
b	1.03	4.20	3.47	-0.69	-5.80
c	-0.34	0.06	6.21	16.49	26.65

An Illustrative Application

Haiti is a suitable country for using the new approach, since marriage patterns are informal. The 1987 Survey of Mortality, Morbidity and Service Utilization in Haiti (EMMUS) included questions on children ever borne, age of mother and age at first pregnancy. Internal patterns in the data suggest that this latter question was answered as age at first birth, and so has been used to tabulate parous women by time since first birth. The 1994 Survey of Mortality and Morbidity in Haiti (EMMUS-II), part of the Demographic and Health Surveys program, collected complete birth histories, providing estimates of child mortality for five year periods back to 15-19 years before the survey. It is therefore possible to apply the technique proposed here to data from the 1987 survey, as well as applying the age-based indirect method, and to compare the resulting estimates to both direct and age-based indirect estimates from the 1994 survey.

Table 4 shows the application of the new method to data on proportions dead tabulated by time since first birth. Coefficients from Tables 2 and 3 for the "West" family of Coale-Demeny model life tables have been used, since the direct birth histories from the 1994 survey suggest a close fit of child mortality patterns to that family. The estimates of $_nq_0$ increase monotonically with time since first birth group, and the $_5q_0$ values implied by each $_nq_0$ in the "West" family of model life tables also increases, except for a very small decline from the 10-14 to the 15-19 years groups. The reference dates of the estimates also increase with time since first birth, from about 2 years before the survey for the 0-4 group to about 12 years before the survey for the 20-24 group.

Figure 2 shows the estimates of $_5q_0$ plotted against reference date, and includes estimates using other methods and data sources. The 1987 EMMUS data have also been analyzed by age group of mother, direct estimates of child mortality by time period are available from the 1994 EMMUS-II, and indirect estimates of child mortality using the age-based methodology have also been made from the 1994 EMMUS-II.

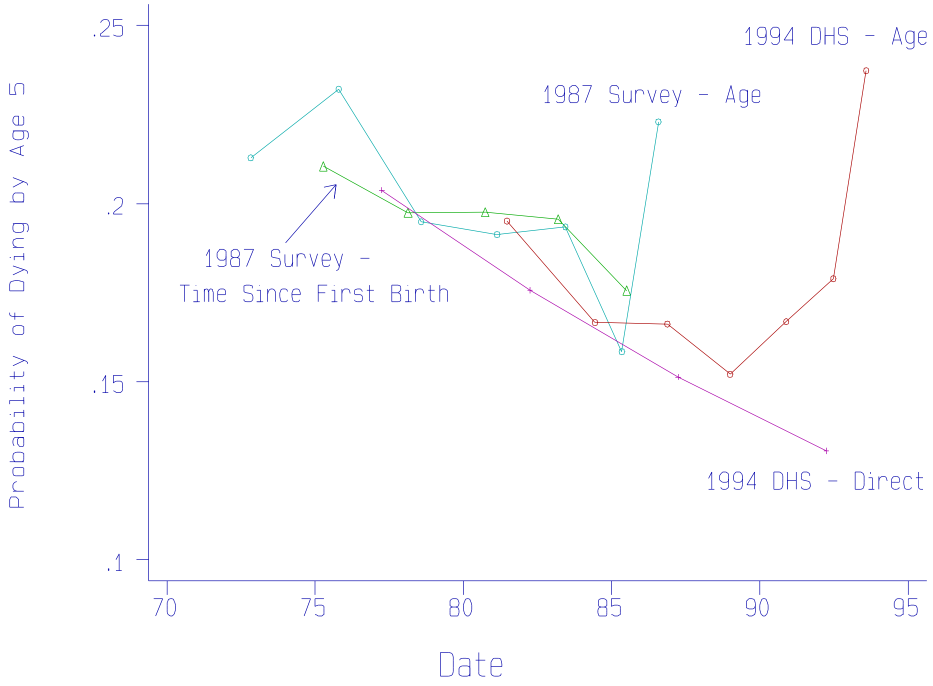
As expected, the age-based indirect estimates, both from the 1987 survey and the 1994 survey, show a sharp upward spike for the estimate nearest the date of the survey (based on women age 15-19). The 1994 direct estimates show a very smooth pattern of decline over time, though at levels generally slightly below the indirect estimates. The new estimates, based on time since first birth, are in general very close to the 1987 indirect estimates based on age group, but the new estimates show a much more consistent time pattern. There is no spike for the most recent estimate, and irregularities observed between the age group estimates based on reports of women aged 35-39, 40-44 and 45-49 are not evident in the time since first birth estimates. The new estimates suggest a period of stalled child mortality decline in the early 1980s that is not evident in the trend shown by the direct estimates from the 1994 EMMUS-II. Both the direct and indirect estimates from EMMUS-II tend to be a bit below the indirect estimates from EMMUS, though the largest difference (excluding mortality estimates based on reports of younger women) between any pair of estimates for a similar time period is less than 25 per thousand.

Table 4: Applications of New Method to Data from Mortality, Morbidity and Service Utilization Survey, Haiti, 1987.

Time Since First Birth Group	Average Children Ever Born	Average Children Dead	Proportion Dead	“West” Family k (i)	n	$_nq_o$	“West” Family $_5q_o$	Time ago	Reference Date
0-4	1.567	0.187	.1193	1.1294	2	.1347	.1755	1.98	85.22
5-9	3.193	0.546	.1710	0.9895	3	.1692	.1955	4.30	83.20
10-14	4.718	0.931	1973	1.0009	5	.1975	.1975	6.77	80.73
15-19	5.754	1.280	.2225	1.0271	10	.2285	.1974	9.36	78.14
20-24	6.216	1.591	.2560	1.0167	15	.2603	.2104	12.22	75.28

$P(1) / P(2) = 0.491$ $P(2) / P(3) = 0.677$

FIGURE 2: Time Since First Birth Estimates of Child Mortality and Other Available Estimates for Haiti: 1970-1994.



Conclusion

Time since first birth provides an alternative proxy for children's exposure to risk of dying. It shares with time since first marriage the advantages over an age based proxy of reduced socio-economic selection combined with large numbers of observations in the shortest exposure category, but avoids the disadvantage of the time since first marriage proxy of not being applicable in populations with substantial proportions of births outside formal unions. Birth rates can be reasonably well modeled by time since first birth, and observed proportions dead can be converted into standard mortality indicators both easily and accurately.

A trial application to data from Haiti confirms the expected advantage of the time since first birth formulation relative to the existing age-based formulation. The new method appears to give good estimates of both levels and trends of child mortality over a 12 year period before the survey.

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