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CHILDHOOD MORTALITY ESTIMATES USING THE *PRECEDING BIRTH* TECHNIQUE: SOME APPLICATIONS AND EXTENSIONS

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INTRODUCTION

The preceding birth technique invented by Brass and Macrae (1985) is being taken up quite widely as a simple method for estimating recent trends in early childhood mortality. Both the questions put to mothers and the analysis of the results are straightforward when mothers are seen at or soon after a subsequent delivery. Several technical aspects of the method which some authors suggested would introduce unanticipated biases into the results, have now been studied and have been shown to be relatively unimportant (Aguirre and Hill, 1987). They include: the effects of a child death on the length of the subsequent birth interval; birth intervals which differ from 30 months; the omission of reports ont he survival of final births or on single children with no siblings. The conclusion from several applications (Brass and Macrae, 1985; Hill and Macrae, 1985; Hill, Traoré, Cluzeau and Thiam, 1986; Arretx, 1984) is that the method appears to deal quite successfully with many of the usual biases introduced by collecting reports on child survival from a non-random sample of women. Although the reports stem from women having a subsequent live birth, the proportion of preceding births dead is a birth-based and not a woman-based measure of mortality. That is, all preceding births are included in the proportion regardless of the age, parity and length of the preceding birth interval. The women contributing reports on the survival of their previous children, although chosen in a non-random way, are largely representative of all women in the general population.

While the method appears to deal effectively with the demographic biases normally associated with the analysis of non-random data, it does not solve the problem introduced when the reports come from an unrepresentative fraction of the total childbearing population. How the mortality of children born to women in clinics compares to the mortality of all children is a matter for separate analysis. The selection effects can work in several ways, as the analysis below illustrates.

It is in part in an attempt to deal with this latter problem of incomplete coverage of the childbearing population that a new initiative has been taken to include the question on the survival of the preceding birth in household surveys. The question remains the same but the analysis is slightly different to allow for the additional complications of additional exposure time and dependence in the survival of pairs of children one born after the other.

I. APPLICATIONS BASED ON THE DATA COLLECTED AT THE TIME OF A SUBSEQUENT BIRTH

There are now several places where the method is being regularly employed to obtain an index of early childhood mortality (see Arretx, 1984). In surroundings in which the great majority of women deliver in hospitals or clinics, and where health services are well developed, few problems are expected. It may be useful for other potential users to summarise experience gained from one systematic trial in five urban maternity clinics (one in a hospital) in Bamako, Mali's capital city. The test brought to light several practical problems and the data illustrate very well the quite subtle biases which can result from selection effects.

a) The setting

In the city of Bamako, some four-fifths of mothers are thought to give birth in clinics or in hospitals. Many women come into the city from outlying areas and of course difficult cases always end up in one of the hospitals. Coverage of all mothers is thus far from complete and an unknown part of the city's child mortality experience described in the clinic and hospital records is "imported" from the surrounding rural areas. In the medium-sized towns, perhaps two-thirds of women give birth in clinics and hospitals although in the rural districts, the vast majority of births occur at home. The rather high proportion of urban women giving birth in clinics is thought to be a tradition carried on from colonial times, when as an attempt to improve child mortality, mothers were required to come to maternity clinics to deliver. The proportions are much lower in anglophone Africa, although in some countries, such as Turkey, home deliveries are followed up with a domestic visit by a health worker. In such circumstances, when most mothers are being visited a few days after a birth, or where the majority of mothers receive a pre-natal visit, the question on the survival of the previous born child could give excellent results.

b) Practical considerations

In every clinic, there is usually some system of recording basic information on the mother and the current delivery. The total number of children born alive and the number still living are frequently collected; in this case, the adapted multiplier technique of Brass and Macrae (1985) or the improved version by M'Backé (1986) could be employed to obtain retrospective estimates of child mortality. In Bamako, these data were available on the existing registers but the questions were not being asked very clearly because of a confusion, possibly quite general, between the terms for a pregnancy or confinement ("geste"), a live birth, a still birth and a miscarriage. "Parity" in medical parlance frequently includes still births and possibly miscarriages too, whereas in the application of the previous birth technique, it is the survival of live born children which concerns us, even if a miscarriage intervenes between the last live born child and the "current" delivery. These points needed careful explanation to the midwives who had been used to adding all live births and still born children together under the heading "parité". This is why in the revised form (Appendix A), the distinction between these different categoriesis carefully drawn. An additional practical point to remember is that we are interested in the survival of <u>all</u> live born children including those who die very soon after birth. The tendency to report very early neo-natal deaths as still-born children is probably quite widespread; the paper-work is usually simpler and it reduces the possibility of the midwife being blamed for poor post-natal care. In both these instances, it is worth stressing that the definitions of terms being used are demographic rather than medical, even though the data are being collected in a clinical context.

In the trial on which further details are available in Hill et al (1986) midwives were asked to complete short supplementary questionnaires including the new questions on the survival of the last (preceding) and second-last births, in addition to completing the usual large maternity registers. There are often vocabulary problems in referring to "last live birth" to "second-last live birth". In Mali, as elsewhere in francophone Africa, the forms and official correspondence are in French whereas the midwives communicated with the mothers in Bambara or another African language. With careful training and close supervision at the outset, these problems can be surmounted. The midwives were used to crossing boxes under selected headings in the old registers; this may be a useful approach where literacy levels are low but a more economic arrangement of the form can be achieved by using code numbers (see Appendix A). Despite initial complaints about the additional work involved, all the midwives filled in the questionnaires accurately and it seemed to us that age and date reporting (the date of the last and second-last births were requested in the trial) were remarkably accurate. There appear to be some advantages in obtaining demographic data in a clinical context compared with doorstep interviewing. Some minor details such as the need to think in advance about the treatment of twins or triplets came to light, but in general all the questions were easily answered by all the mothers, including those who only spent a few hours in the clinic. All the questions except those on the current delivery can be put to the mother at her first entry into the clinic.

Although the numbers of previous born children alive and dead can be quickly obtained by simple addition of the appropriate columns in the registers, it seems very likely that some further analysis will be required, using the other variables such as maternal age, parity and birth weight of the new-born. Re-copying all these variables for some hundreds or thousands of deliveries seems an unattractive option. One possibility is to use carbonised forms with the carbon copies being regularly dispatched for data processing. One newer and very appealing option is to use hand-held computers for data capture. In this case, the results could be entered once by an itinerant data collector, verified by the program in the machine, and then transferred to a central micro- or larger computer for speedy analysis and feedback of information to the clinics.

c) Analysis and results

The basic calculation of the index of early childhood mortality, approximately q(2), from the total number of previous born children alive and dead is trivial. Equally straightforward is the calculation of an estimate of q(5) from the number of second-last children alive and dead (Table 1). The interesting part is the comparison of the indices with each other and with other sources, and the further disaggregation of the results.

The proportion of previous children dead amongst all 4775 births recorded during the trial period (January-April 1985) was 0.142, a figure which translates into an infant mortality rate of about 100 per thousand using the African standard model life table. In general, the figure matches quite closely the results obtained by retrospective questioning of mothers in a large survey conducted amongst a representative sample of Bamako households by the Sahel Institute at about the same time.

In the Sahel Institute's multi-round follow-up survey of a sample of births occurring in Bamako's hospitals and clinics, the unadjusted infant mortality rate was estimated to be 63 per 1000, but it is widely recognised that this figure is too low because of severe loss to follow-up between the birth and the first home visit (Dicko 1986). In the clinic study, the proportions of second-last births dead indicated an infant mortality rate of about 90 per 1000, lower than the figure derived from the proportions of last-born children dead. Although some under-reporting of children who died some time before the current birth cannot be excluded, a more likely explanation is that we are looking at the effects of selection in the results.

One indication of these effects can be seen when we compare the childhood mortality estimates from the survival of the two previous born children with those obtained by application of Brass and Macrae's adapted multiplying factor technique to the proportions of all previous born children dead at the most recent delivery. The calculations are shown in Tables 2 and 3 and the resulting mortality estimates are plotted in Figure 1. The discordance between the two series is striking and both suggest a rather unlikely worsening of child mortality in the period before the enquiry. The difference in level between the two seems to us to be attributable to the effects of selection which appear to be related to the age of the mothers.

We can see how these effects might be working from two tables of the proportions of previous children dead by the age and parity of the mothers. In Table 4, we see a steady and highly implausible fall in the proportions dead by age; the trend is less clear by parity but the proportions dead beyond parity 5 do not show the expected trend (Table 5)

Some further examination of the characteristics of the mothers giving birth in clinics by educational status and comparison with the composition of the general population of mothers as revealed in the Bamako household survey (see earlier) revealed that proportionately more younger women were giving birth in clinics and proportionately fewer older women. Many socio-economic and demographic characteristics (education, income, parity etc.) are age-related so that on average, the older women seen in the clinics were from the relatively higher social classes. It seems logical that the children of these women would have lower than average mortality than children of lower class women of the same age. This greater preference of upper class women to give birth in hospitals and clinics would explain the rather unlikely trends in childhood mortality indicated in Figure 1 and in Tables 4 and 5. The discordance between the results from the two methods is probably attributable to selection effects as well since we can see from Table 2, last column, that the adjusted proportions of all children dead by age do not rise steeply enough, particularly beyond age 30.

Several general points are worth noting by way of conclusion. First, the pattern of selection in situations where not all mothers give birth in clinics is likely to vary from one place to another. In Bamako, it seems that better educated younger women are over-represented as clinic-attenders, a pattern which may be quite widespread. One could imagine quite different patterns however, if the basis for selection was different, e.g. only women with previous obstetric problems or difficulties with the current delivery came to clinics. Secondly, if the pattern of selection is to over-represent the upper class mothers, then the resulting childhood mortality estimates must be treated as minimum estimates. It seems to us unlikely that infant mortality in Bamako is any less than 100 per 1000. Thirdly, although the stress in this analysis has been on the production of life table measures of child mortality, the main use foreseen for the previous born child technique is as an index for tracking changes in level of early childhood mortality. With a regular system of data collection, the principal value of the results would be to provide a month by month (where numbers permit) plot of the index of early childhood mortality. Used in this way, many of the selection problems discussed earlier will be much less important. Finally, it is worth stressing the gains to be obtained from collecting data on a routine basis as part of the health information system. We have already referred to the accuracy and completeness of the answers provided by mothers in the pilot study. In addition, extra variables can be readily obtained, such as the valuable information on birth weights, (see Table 6).

The cross-tabulation of the birth weight of the new-born against the survival of the two previous children speaks volumes about the concentration of the excess child mortality risks in certain families. It might be possible to retrieve the birth weight of the previous born child in certain circumstances e.g. where the mother previously gave birth in a clinic and can provide an approximate date of birth. Birth weight may also be recorded on patient-held cards. Clearly, there are many possible ways in which the basic application described above might be extended to other situations. An important point not to be overlooked is the low cost of the whole exercise. All the data collection was carried out by the midwives already in place in the clinics.

II. ADJUSTING THE PROPORTIONS OF PRECEDING CHILDREN DEAD WHEN MOTHERS ARE INTERVIEWED AT TIMES OTHER THAN AT A SUBSEQUENT DELIVERY

(a) Some possible applications

One of the inevitable difficulties in estimating early childhood mortality from information provided by women giving birth in clinics is that the estimate only applies to that part of population giving birth in clinics or hospitals. Of course, if all women give birth in maternity hospitals, then the estimate of child mortality applies to the whole population, but this is not likely to be the case for Africa, at least for some time in the future. To improve the coverage of the reports, one option is to make the questions on the survival of the previous born child part of the routine reporting which, in an effective primary health care scheme, village health workers (or whatever the lowest level in the hierachy of health workers is called) would provide. At the time a new birth is reported to the "key" village or health centre, the mother could be asked about the survival of the previous born child (if any). Even for the illiterate village health worker, a simple tick or cross reporting form could be developed. Another possibility is to pose the question to pregnant women who may be seen at some time before the birth. If more of the female population are seen in ante-natal clinics than in maternity clinics (i.e. most deliver at home), then the resulting mortality estimate will be more representative of the general population. Some care will be needed to avoid double counting the same women in the two recording systems. Reports on the proportions of previous children dead obtained from pregnant women will be indistinguishable from those obtained at the time of the delivery since a reduction in the duration of the period of exposure of the previous born child to the risks of dying by a month or two is unimportant. Mortality risks around the second birthday change only slightly within the space of a few months, as Table 7 clearly illustrates.

A much more interesting idea is to collect the information of the survival of the last born child during the course of an immunisation programme, an immunization coverage survey or another similar health intervention which is designed to reach all the young children in an area. In these circumstances, the problem of selection bias associated with using data from clinic attenders disappears, but a new source of bias is introduced. If the immunisation programme, say, is aimed at young children, only mothers with surviving young children will be contacted. There is an association between the survival of the last born and previous born child in every population because excess mortality risks tend to be concentrated in some sub-sections of the population - poor, uneducated, badly housed, low-class families; the data from the Mali clinics reported in Table 6 earlier are a good illustration of this. The strength of this association between the survival of pairs of successive children can be estimated, however, and as we show later, it varies systematically with the overall level of child mortality in the population and the age of the last born children at interview. Thus, adjustment factors can be developed so that reports on the survival of the previous born child obtained only from mothers with a surviving child born subsequently can be corrected to bring them into line with the same information obtained at the time of the delivery. These adjustments must also take into consideration the additional time the previous born child has been exposed to the risks of dying since this child will have been exposed for a period consisting of I, the mean birth interval, plus y, the age of the surviving last born child. In fact, the effect of this additional y months of exposure on the proportions of preceding children dead is not very marked for the reason given earlier; the risk of dying between say age 2 and age 2½ or even 3 years are not changing as rapidly as during the first 6 or 12 months of life.

(b) Calculating the association between the survival of two successive births

The problem of calculating the degree of association between the survival of two successive children can be reduced to a very simple form by setting out the various combinations of possible outcomes in a two-by-two table (Table 8). Assume for the moment that all previous born children (the older) were born x months ago and all the last born children (the younger) were born y months ago. All the older children are assumed to be the same age X, and all the younger children, the same age Y. The probability that the older child in a pair is dead is:

$$\Pr\text{ (older is dead)} = \frac{B + D}{A+B+C+D} = q\ (x)$$

C and D are unknowns, however, since when the younger child is dead, the mother is not interviewed, but an estimate of q (x), $\hat{q}(x)$, can be obtained as:

$$\hat{q}\ (x) = \frac{B}{A + B}$$

the conditional probability that the older child is dead when the younger child is alive. The other conditional probability for which no information exists is:

$$q^{c}(x) = \frac{D}{C + D}$$

the probability that the older child is dead when the younger child is also dead. The true probability q(x), is a weighted average of the last two conditional probabilities. The same logic applies to the conditional probabilities for the younger child:

$$\hat{q}(y) = \frac{C}{A + C} = \text{Pr (younger child dead/older alive)}$$

$$q^c(y) = \frac{D}{B + D} = \text{Pr (younger child dead/older dead)}$$

When we have two sets of conditional probabilities like these and we assume the probabilities are independent, we can calculate:

$$\text{Pr (older child dead and younger alive)} = q(x)\left[1 - q(y)\right]$$

$$\text{then } \hat{q}(x) = \frac{q(x)\left[1 - q(y)\right]}{1 - q(y)} = q(x)$$

But the most plausible situation is not independence of the two outcomes, but a positive dependence between pairs of successive children i.e. the alive-alive and dead-dead combinations are the most likely although we cannot exclude the possibility of negative dependence which might arise say, if a mother with a preceding child dead was so carefully monitored that the chances of her succeeding child surviving were raised much above the average. Let the factor of dependence be f in such a way that:

$$q^c(y) = \text{Pr (younger child dead/older dead)} = f.q(y)$$

As the degree of dependence increases, so does f. The question then is how good an estimation of $q(x)$ is $\hat{q}(x)$ with different values of f? The answer to this question can be found in Table 10 where the percentage underestimation of the index of early childhood mortality is shown for various values of f (the rows) and selected mortality levels (the columns). For simplicity, the mean age of the last born children has been set to one year exactly. The percentage underestimation (PCU) is:

$$\text{PCU}\left[\hat{q}(x)\right] = 100\,\frac{q(x) - \hat{q}(x)}{q(x)} = 100\,\frac{q(x) - \dfrac{q(x)\left[1 - f.q(y)\right]}{1 - q(y)}}{q(x)}$$

$$= 100\,\left[f - 1\right]\frac{q(y)}{1 - q(y)}$$

The formula for the correction factors is:

$$F = \frac{q(x)}{\dot{q}(x)} = \frac{q(x)}{\frac{q(x)\,[1 - f\,q(y)]}{1 - q(y)}} = \frac{1 - q(y)}{1 - f\,q(y)}$$

A selection of these percentages of underestimation is shown in Table 10 and adjustment factors derived from this formula are shown in Table 11.

As Table 10 indicates, the percentage underestimation of the index of early childhood mortality rises with increasing values for f, the degree of association between the survival of successive pairs of children, and with the mortality level. With a modest degree of dependence, say f = 2, and on infant mortality level of say 150 per 1000, the percentage underestimation is 17.65 per cent. Thus, it seems worthwhile to consider the development of a set of adjustment factors for a variety of circumstances in which we encounter different values of f and different levels of early child mortality. There are two separate problems to consider. One is to calculate some empirical values for f in order to provide some guidance for users in choosing a likely value for f in their own particular population.

Secondly, we have to consider the effect on the index of early childhood mortality of the additional period of exposure of the previous born child to the risks of dying when the data on the proportions of these children dead are collected from mothers with a subsequent living child aged 6, 12, 18 etc. months at interview. Clearly, the previous born child has been exposed for a period I, the mean birth interval, plus an additional period y where y is the age of the last born surviving child.

(c) Empirical values for f, the factor of dependence

The most accessible source for the calculation of the degree of association between the survival of successive pairs of children is the set of maternity histories collected for some forty-two countries by the World Fertility Survey. Since we know that the degree of dependence varies with the level of mortality, we have selected just five countries with different levels of child mortality for this test. The f values were calculated directly by examining systematically the survival status of pairs of children in each birth history record. Women with only one birth are of course excluded. By moving month by month through each birth history, we are able to calculate the value of f for different ages of the second-born child. The results are summarised in Figure 2.

There we observe the predicted relationship between the f's and the mortality level. In brief, it seems that at higher levels of mortality, risks of a child dying are distributed more evenly throughout the population. When mortality is lower, the risks of dying are less randomly distributed and instead are more concentrated amongst certain groups of high risk mothers such as those who have already experienced a child loss. The interesting conclusion for our present purpose is that the values of f when the most recent born child is aged six months or over are quite close to two for the mortality levels found in many developing countries (i.e. infant mortality rates of around 100 per 1000). Thus, in our search for adjustment factors to apply to the proportions of previous born children dead obtained when a mother is seen with a subsequent living child aged x months, we can assume an initial value of f = 2 when f is unknown for most applications in populations with moderate to high child mortality. A higher value might be assumed when mortality is lower, as in the Guyana example on Figure 2. As Table 10 and the more detailed tables in Aguirre and Hill (1987) show, a slightly wrong assumption about the true value of f will not greatly affect the adjustment factors on the revised value of the index of early childhood mortality. With f = 2, and an estimated infant mortality rate of about 100 per 1000, the reported proportions of previous children dead will be raised by a factor of 1.134 when the last born children are one year old at interview.

(d) The effect of additional months of exposure

When the reports on the survival of the previous born child are obtained not at the time of a subsequent birth but some y months after, the previous born child will have been exposed to the risks of dying for the length of the mean birth interval, I, plus an additional period of y months, the age of the last born child at interview. All other things being equal, this extra exposure will increase the proportions of previous born children dead with reference to the case when the reports are obtained from mothers giving birth on a subsequent occasion. This upward bias in the proportions is independent of the correspondence between the survival of pairs of successive children discussed above. How large are the effects of this additional period of exposure on the proportions of previous born children dead?

As indicated earlier, the proportions of previous children dead change relatively slowly with small additional periods of exposure beyond two years. This is because the life table function q(x) is beginning to flatten out around age two; this is true of any level of mortality. As Table 7 illustrates, q(2) is in most cases around 90 per cent of the value of q(3). Whether a further adjustment to the proportions of previous children dead to take account of small additional periods of exposure is justified depends on several factors. One is of course the length of this additional period. If the mean age of the subsequently born children is one year, then based on the results in Table 7, a 10 per cent downward adjustment may be applied. In most applications, it seems to us that there is a case for not using the data from groups of mothers whose last born living children are much over 18 months old, simply to minimise omissions and other problems of recall. A second consideration in examining the case for a further adjustment is whether the original reports are really of a high enough quality to justify an adjustment as subtle as 10 per cent. Finally, an important consideration is whether there is a need at all to correct the data collected in this way. If the aim is simply to produce an index of early childhood mortality and the population reporting will remain the same (i.e. the ages of the last born living children are about the same, as they might be, say in a vaccination programme), then no adjustments for either the association between the survival of successive pairs of children or for the extra period of exposure of the previous born child are really necessary. In this case, changes in the index are of primary importance and translation of the index into a life table measure of child mortality is unnecessary. It may be, however, that data from different sources are to be compared. For example, it may be desirable to make comparisons between the proportions of previous born children dead from reports obtained at the moment of a subsequent birth with those obtained from mothers with a subsequent living child aged y months when interviewed. In this case, both the adjustments for dependence and the additional period of exposure would be necessary, although as we shall see, these adjustments work in opposite directions and may indeed cancel each other out in some circumstances. For example, take the case where the estimated infant mortality rate is about 100 per 1000, f = 2, and the age of the last born children is one year. The correction factor from Table 10 will be 1.13. Assume that the proportion of previous children dead at interview is 0.2. Then the adjusted proportion dead will be:

$$0.2 \times 1.13 \times 0.9 = 0.203$$

an insignificant change in the original value.

III. THE PRECEDING BIRTH TECHNIQUE APPLIED IN HOUSEHOLD SURVEYS: THE TRIAL IN LIMA, PERU 1987

The trial was conducted in marginal areas located in the outskirts of Lima. These settlements are relatively new. They have developed mainly in the last two decades. For this reason they are called 'pueblos jovenes' (young towns). Most of the recent rapid population growth of Metropolitan Lima has been accommodated by this form of unofficial, often illegal urban expansion. People living in the 'pueblos jovenes' either come from other older deprived residential areas of Lima or are immigrants from the highlands. The growth of the 'pueblos jovenes' has been so rapid that the most recent available vital statistics classified by district are out-dated since they come from a time when the configuration of the 'pueblos jovenes' was rather different. These statistics thus cannot be used to measure recent childhood mortality levels. The main of the study was to test two short-cut methods for obtaining recent estimates of early childhood mortality in countries with inadequate vital registration. The idea is that a few simple supplementary questions added to routine household health surveys (EPI or CDD surveys for example), coupled with some novel forms of analysis, might produce improved childhood mortality estimates. Hence the study in Peru included several extra questions beyond the minimum set so that the mortality estimates arrived at by different methods could be compared. In essence, the main goal was to be able to say what questions in future studies could be confidently omitted from simple household surveys in order to reduce the size and complexity of both the questionnaire and the analysis.

The mortality estimates here are based on two separate sources of information in the questionnaire. First, all eligible women were asked about the total number of live births and the number of their children alive at interview. The proportions dead of children ever-borne are converted into life table measures of child survival and approximate time references for each of the estimates calculated. The key element is always the quality of the basic data. The drawbacks of the method for health workers include:

* the unreliability of, the most recent estimates of mortality based on reports from women aged 15-19 and 20-24. Usually, the mortality levels estimated from these women are too high because of selection effects, higher than average mortality of first born, and small numbers.
* the difficulty of producing mortality estimates for precise periods because the dates of birth for each child are not collected. Models are used to re-distribute births over the pre-survey period.
* the large size of the sample required to calculate Brass-type estimates of childhood mortality.

Secondly, in this study, each woman, in addition to the questions on the total number of children ever-borne and surviving, was asked to provide a brief birth history for her last three live births. The dates of birth and dates of death (where applicable) of each child were collected in months and years. This is a variation of the 'truncated birth history' approach which has been used in several large household surveys including some conducted recently by the Westinghouse Demographic and Health Surveys (e.g. in Ondo State, Nigeria) and in experimental surveys in Peru and the Dominican Republic (see Goldman, Westoff and Moreno Navarro for details). In these kinds of surveys, an attempt was made to collect all births within the five-year reference period preceding the interview. Experiments have shown that there are major problems with this approach for fertility measurement because of erroneous dating of births (Potter, 1977). For mortality studies these effects may be less important but in any case, the focus on the last three births regardless of when they occurred prevents interviewers reducing their work by displacing births to a date outside a specified reference period.

These truncated birth history data have been analyzed in two ways. First, monthly life tables have been constructed for all births occurring in specified periods before the survey. This produces mortality estimates for a series of birth cohorts in the pre-survey period. This form of analysis is vulnerable to quite small omissions of dead children, particularly those who died soon after birth. In addition, there may be some selection biases if there is a strong link between childhood mortality and parity.

A second quite different form of analysis, one which is relatively new and at the heart of this trial, can also be performed on the same data. Using a variation of the preceding birth technique, we calculate directly the proportion of second-to-last children dead at the time of interview. When the women are interviewed at the time of a birth, this proportion is quite close to $_2q_0$, the probability of dying between birth and the second birthday. When women are interviewed soon after the last birth regardless of the survival status of this child, as was the case in Peru, the proportions of preceding children dead estimates survival from birth to some higher age. This age is approximately 4/5 the length of the mean birth interval, say 4/5 x 30 months, plus the length of the period between the most recent birth and the date of the interview. By restricting the calculations to women with last born children born in fixed periods before the survey such as in the last two years, probabilities of dying can be calculated and dated.

There are many technical issues connected with the selection effects introduced by interviewing women with a recently born child. These are currently being explored in more detail and will be the subject of a separate report.

An essential feature of the trial was to be able to compare the childhood mortality estimates derived from the two separate data sources calculated by alternative methods. The mortality indices describe the probability of survival from birth to a variety of different ages. A standard measure is required for comparison. The most convenient for our purposes is the level parameter, alpha, from Brass's relational model life table system (Brass, 1971). We use the General Standard life table throughout as the reference standard.

MORTALITY RESULTS

a) The Brass indirect estimates

From the total numbers of children ever-borne and surviving, we calculate the life table probabilities of dying by age (x) for the 15 or so years before the survey. For comparison of the results for women of different ages and to facilitate comparison with estimates from other sources, all the $_xq_0$ values have been converted to $_2q_0$ values by interpolation in the logits of Brass's General Standard lifé table. The results, together with the alpha values which are the level parameters in the logit system and the time location of each mortality estimate are shown in Table 11. The general pattern of the results can be seen more clearly from Figure 3. The reports from women aged 25-49 years at interview indicate that a steady improvement in childhood mortality has been underway over the whole 15 year period before the survey even amongst the current residents of the 'pueblos jovenes' in the sample. This improvement has been quite general in Peru, as Moser showed in her analysis of national-level census and survey data (Moser, 1985:16-22). The data from women aged 15-19 and 20-24 at survey are misleading if taken at face value. Here, as in all five of the censuses and in the 1977 survey, the mortality levels for these younger women cannot be regarded as typical of the level of mortality prevailing amongst children of all mothers. Selection effects related to age and parity are known to strongly influence the results from these two youngest age groups of women. Since the trend in the alpha values for the period 4 to 15 years before the survey is so consistent, a straight line has been fitted to the five points for women aged 25-49 by ordinary least squares regression (r^2 = 0.95) and the two most recent alpha values estimated by extrapolation of this trend line. The results are shown in column 7 of Table 11. There, the probabilities of dying before age 2 in column 8, $_2q_0$, form a quite consistent series with the most recent extrapolated values around 49-50 per 1000 for the period 1 to 3 years before the survey. In the General Standard life table, these values correspond to a infant mortality rate of between 36 and 38 per 1000. In general, the Brass method after the adjustment of the data for the two youngest age groups gives a good idea of the overall trend in childhood mortality in the pre-survey period. The method cannot isolate year to year fluctuations in mortality nor can it pick up changes in the age pattern of death. It does warn us of the major improvements in child survival which have been taking place in Peru including the poorer quarters of Lima which complicate the task of estimating early childhood mortality for the period just before the survey. In addition, the low level of infant and childhood mortality is remarkable, given that those districts were supposed to contain the urban poor. With such low mortality, there is a problem of locating enough childhood deaths for some forms of analysis, particularly for sub-groups in the survey.

b) Life table estimates

The next step is the comparison of these indirect estimates with the life table estimates for the same time periods using the supplementary information on the dates of birth and dates of death contained in the partial birth histories. As Table 12, section B shows, the life table estimate of $_2q_0$ for all the children born in the five years before the survey is 48 per 1000, almost identical to the adjusted Brass estimate of 50 per 1000 for the same period. The infant mortality rate from the life table is 40 per 1000, 9 points lower than the equivalent adjusted Brass estimate. The Brass estimate of infant mortality, however, is based on an extrapolation of the trend over the period 4-15 years before the survey and is therefore less reliable for dates outside this period.

No great significance can be attached to the difference between the life table and the adjusted Brass estimate of infant mortality although the comparison of the $_2q_0$'s is quite soundly based. The two estimates of mortality before age 2, one from the Brass technique, the other from the life table for all children born less than 5 years before the survey, are effectively the same once allowance is made for sampling variations.

More detailed life tables can be calculated from the data on the three last born children. Values of $_2q_0$ for both two year and five year time periods are shown in Table 13. There is some irregularity in the two-year series largely connected with reporting errors. It appears that some omission of dead children has occurred for the most recent period, possibly because of the reluctance noted above of some women to start their birth history with a report on a dead child. In addition, the apparent worsening of mortality for the period before 1980 is highly implausible and is probably linked to some omission of dead children. Selection effects are also important because the women reporting on their third-to-last live birth are necessarily older and will on average have a higher parity than all women who give birth in a set period regardless of age and birth order. This is one of the drawbacks of collecting only partial birth histories and analyzing them in this way. By selecting births of a particular order for analysis, one is in effect selecting women and not births because each woman is being allowed to contribute only one birth to the analysis. For the study of childhood mortality, of course, it is rates based on births and not on women which is at the centre of our concern.

c). Previous born child technique

Apart from its simplicity, particularly in the form used when mothers are interviewed close to the time of a delivery, the previous born child technique is an attractive method for estimating early childhood mortality because it takes care of most of the complex selection effects when estimates of child mortality are mother-based. With the previous born child technique, although the reports come only from women with a recent birth, women of all ages and parities can contribute to the index. In addition, the women are not selected with reference to birth interval length because no reference period for the date of the preceding birth is specified. Apart from the well-known biases resulting from covering only the women seen at the time of delivery, the main systematic selection effect, that related to the omission of reports on the survival of the last child in the family, is thought to be unimportant (Aguirre & Hill, 1987).

In the Peru trial, of course, women were not interviewed at the time of a delivery but at various times after their last live birth. Since the time elapsed since this last birth is known directly from the truncated birth histories, we can calculate the proportion of previous children dead for women who have had a recent, subsequent birth. Thus, amongst women reporting a live birth in the 24 month period before the survey, 57.8 per 1000 of their previous born children were dead at interview. In this case, we can calculate directly the mean length of the last closed birth interval (the interval between the previous and the most recent born children = 41.4 months). We also know that the dates of birth of the last born children born in the 24 month period before the survey are centred 10.4 months before the interview. Hence, the proportion of preceding children dead amongst women with a subsequent live birth in the last 24 months (0.058) is not the probability of dying by age two but of dying by some higher age. This age is approximately 0.8 x 41.4 + 10.4 months = 43.5 months or 3.6 years. By interpolation in the logits of Brass General Standard life table, we can find the value of ${}_2q_0$ corresponding to a ${}_{3.6}q_0$ value of 0.058. The result is 0.0496, equivalent to an alpha value of -0.761. This estimate refers to mortality conditions prevailing 3.1 years before the survey, as calculated from the data on date of death.

The same logic can be applied to the proportion of second-to-last children dead at the time of the interview, although here selection effects become more important as the proportion is based only on reports from women with at least three live births. The proportion of these second-to-last children dead is 0.075. The mean length of the next-to-last closed birth interval is 34 months. Hence, the proportion 0.075 is approximately ${}_xq_0$ where x = 5.42 years by direct calculation. Interpolation in the General Standard results in an alpha value of -0.673, and the equivalent value of ${}_2q_0$ is estimated as 0.0542. This refers to a period centred about 3.1 years before the survey because of the concentration of child deaths at the early ages. These results are summarised in section C of Table 12.

All the mortality estimates arrived at by these different methods are presented in Table 12. Since each estimate is located in a particular time period and childhood mortality has been improving steadily in the 15 years before the survey, direct comparisons are easier when the alpha values are plotted on a graph. In Figure 3, we see that the alpha values which describe the level of mortality in early childhood are very close for all three methods - the Brass indirect estimates, the life table for births in the last five years, and the preceding birth technique results. Unfortunately, the true level of mortality in this population is unknown so absolute comparisons are impossible, but the similarity of results from the adjusted Brass estimates to the life table results and the previous born child technique is most encouraging. The slightly higher estimate of mortality obtained from the proportions of second-to-last children dead may be simply due to small numbers (680 births); but the selection effects referred to above may also be responsible.

DISCUSSION OF FINDINGS

In general, the interviews took only a little longer than the basic EPI coverage survey. The idea of adding these mortality questions to other health surveys therefore seems perfectly reasonable.

Age, birth order and birth interval length are just three main demographic factors which strongly influence child survival. There are other equally strong socio-economic factors such as maternal education and household income. In the case of the demographic factors in particular, care must be exercised in order to avoid inadvertantly selecting a biased sample of mothers for interview simply because of the way eligible respondents are identified. Thus, calculation of life tables for recently born children by order (last born, second-to-last born etc) is a bad way to measure recent childhood mortality levels as it precludes mothers from contributing more than one birth to the calculations. What is required is a birth-based sample to which mothers of all ages, parities, birth intervals etc, can contribute in roughly the proportions found in the general population.

Little more needs to be said here about the well-established Brass technique for measuring childhood mortality. The preceding birth technique is new, particularly in the form it was used in the Peru trial. When restricted to mothers with a recently born subsequent child, the proportion of preceding born children dead at the time of interview can be used to produce recent estimates of ${}_{2}q^{2}$ which closely match the life table measures for all births in the preceding 5 years and the adjusted Brass estimates (see Figure 3). The disturbing finding is that to obtain about 1000 preceding births born to mothers with a last birth in the 2 years before the survey required the interviewing of 7500 women of reproductive age. The problem arises from the need to restrict the reference period for the occurrence of the subsequent birth. One solution would be to screen women for interview with a preliminary question on the date of their last live birth but the difficulty here would be omission of some last born children who had died by the time of the interview. If such omissions appear to be serious, the analysis could be based entirely on the reports of previous children dead obtained from women with a surviving subsequent child. In this case, the adjustment for dependence in the survivorship of pairs of children, the preceding and the most recent children, would need to be applied. The Peru data are not a good test of this method because of the small number of previous born children dead. In fact, the proportion of preceding children dead to women whose most recent birth (born in the 24 months before the survey) was alive at interview is almost the same as for all women. The figures are:

* Most recent birth alive, % preceding birth dead = 60/1022 = 0.59
* All most recent births, % preceding births dead = 61/1056 = 0.58

Even including all the most recent births in the last five years in order to obtain more deaths, the proportions are also quite similar:

* Most recent birth alive, % preceding birth dead = 107/2050 = 0.52
* All most recent births, % preceding births dead = 112/2128 = 0.53

The main conclusion from these results is that mothers whose last born child is dead are omitting to report this information in some of the interviews. The effects of this omission are not serious in the case of Peru when mortality is being estimated by the preceding birth technique because in a population such as this, the number of women with last born dead children is quite small. Application of the adjustment proposed by Aguirre and Hill provides an idea of the scale of the effects of omission. Assume all the most recent children born in the 24 months have an average age of one year at interview. Given that infant mortality is close to 50 per 1000, the factor of dependence between the survival of successive born siblings can be estimated to about 2.5. Using the tables provided in Appendix C by Aguirre and Hill, we find that the proportion of preceding children dead reported by women whose most recent child is alive has to be raised by about 1.087 to produce the corrected proportion dead. This is a small adjustment which would raise the estimate of the infant mortality rate, for example, by about 3 points per 1000. With infant mortality rates closer to 150 per 1000, the adjustment becomes more significant as the correction factor will be closer to 1.21 in this case.

IV. RECOMMENDATIONS FOR FUTURE SURVEYS

The data produced by this trial have helped to clarify several important issues for the design of future surveys meant to measure recent levels of childhood mortality in populations without full registration of births and deaths. Although the choice of Lima's 'pueblos jovenes' was a good one as far as the conduct of the survey and the quality of the age and date reporting was concerned, the low level of mortality led to the collection of smaller numbers of deaths than anticipated. The ideal site for a second trial would be somewhere with high mortality and a reliable independent source of information on child survival. A 'population laboratory' or a site with an established demographic surveillance system would be ideal.

Leaving aside the consideration of additional trials, what conclusions can be drawn for the design of future childhood mortality measurement surveys so far?

a) The principle of collecting childhood mortality data in brief household interviews is not in question. Enough women can be interviewed with a team of less than 20 people in about 3 weeks to yield large enough numbers for most forms of analysis. The costs and infrastructure required for such a survey are not very different from the needs of the average national EPI coverage survey.

b) As far as possible all women likely to bear children should be interviewed. Some care is needed not to omit women who are difficult to locate so some call-backs must be included in the survey protocol. Only some forms of proxy reporting is acceptable.

c) The minimum set of questions on childhood mortality include:

* the Brass questions on children ever-borne and surviving
* the dates of birth and the dates of death (where applicable) of the two most recent live births

d) These data must be analyzed on a 'per birth' basis; that is, use the Brass method on the data on children ever-borne and surviving, remembering that the parity data for calculating the adjustments must refer to all women and not just the ever-married. Life tables can be calculated for particular reference periods but should not be calculated for the last or second-to-last births separately. The preceding birth technique can be used provided the data are for women whose last birth was quite recent, say within two to three years of the survey. The calculations may be performed separately on the data from women whose last live birth is alive at interview. Adjustments for the effects of dependence may then be necessary if omissions of last born dead children are substantial and where mortality is high.

e) The restriction of the sample to younger women is worth considering if reduction of field work costs is important. Leaving out women over age 35 will not greatly affect the results in most circumstances. Omission of teenage mothers would introduce some serious biases because of the excess mortality usually experienced by their children.

f) Although the trial in Peru and others currently underway or planned in Jordan or Zaire show that some additional questions on childhood mortality can produce some more upto date estimates than older methods, the real strength of the preceding birth technique is when it is applied to regularly collected information. In such circumstances, time trends should be much easier to detect than in retrospective survey data. This is the main focus of interest of the major health agencies.

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TABLE 1 CHILD MORTALITY ESTIMATED FROM PROPORTIONS DEAD OF LAST AND SECOND-TO-LAST BIRTHS

	LAST BIRTHS	SECOND-TO-LAST BIRTHS
Births	4775	3737
Deaths	679	620
Proportions Dead	0.142	0.166
Probabilities of dying		
${}_1q_0$	0.101	.091
${}_2q_0$	(0.142)	.128
${}_5q_0$	0.184	(.166)

Notes:

1. The numbers in brackets are the observed figures
2. The probabilities of dying were interpolated from the Brass African Standard Life Table (β = 1)

TABLE 2 ADJUSTMENT OF REPORTED NUMBERS OF CHILDREN EVERBORN OBTAINED AT THE TIME OF A DELIVERY TO PRODUCE REVISED PROPORTIONS OF CHILDREN DEAD BY AGE OF MOTHER

AGE OF MOTHER	Number of women (F)	Total live births (NV)	NV + ½F	NV + 0.2F	PARITY	TOTAL DEAD CHILDREN	%DEAD
15-19	1353	444	1120	715	-.828	103	14.4
20-24	1560	2390	3170	2702	2.032	504	18.6
25-29	1636	5244	6062	5571	3.705	1110	19.9
30-34	965	4898	5380	5091	5.575	1118	22.0
35-39	536	3805	4073	3912	7.599	834	21.3
40-44	111	954	1010	976	9.099	199	20.4

TABLE 3 PROBABILITIES OF DYING ESTIMATED FROM THE REVISED PROPORTIONS DEAD BY AGE OF MOTHER

i	AGE OF MOTHER	$_iq_o$	YEARS BEFORE THE SURVEY t	INTERPOLATED VALUES $_1q_o$	$_2q_0$	$_5q_0$
1	15-19	.137	1.3	(.137)	.189	.241
2	20-24	.194	2.5	.141	(.194)	.247
3	25-29	.203	4.1	.129	.178	.228
5	30-34	.227	6.0	.128	.178	(.227)
10	35-39	.225	8.1	.106	.148	.191

Notes:

1. The parities for married women on Table 3 have been multiplied by the proportions of married women in the district of Bamako according to the Census of 1976 in order to calculate the parities for all women.

i.e. $P_1 = 0.828 \times .304 = 0.252$

$P_2 = 2.032 \times .650 = 1.322$

$P_3 = 3.705 \times .835 = 3.094$

$\therefore \quad P_1/P_2 = 0.190$ et $P_2/P_3 = 0.427$

2. The probabilities of dying were calculated using Brass's African standard model life table. The figures in brackets are the original results.

TABLE 4 CHILD MORTALITY ESTIMATED FROM PROPORTIONS DEAD OF LAST AND SECOND-TO-LAST BIRTHS BY MOTHER'S AGE

A. LAST BIRTHS

AGE OF MOTHER	BIRTHS	DEATHS	PROPORTIONS DEAD ($_2q_0$)
15-19	362	79	0.218
20-24	1248	198	0.159
25-29	1549	213	0.138
30-34	949	118	0.124
35-39	529	55	0.104
40-44	111	11	0.099
45+	15	2	0.133

B. SECOND-TO-LAST BIRTHS

AGE OF MOTHER	BIRTHS	DEATHS	PROPORTIONS DEAD ($_2q_0$)
15-19	59	18	0.305
20-24	697	163	0.234
25-29	1394	238	0.171
30-34	929	125	0.135
35-39	525	57	0.109
40-44	110	16	0.145
45+	15	2	0.133

TABLE 5 CHILD MORTALITY ESTIMATED FROM PROPORTIONS DEAD OF LAST AND SECOND-TO-LAST BIRTHS BY MOTHER'S PARITY

PARITY	N	PROPORTIONS DEAD OF LAST BIRTHS ($_2q_0$)	N	PROPORTIONS DEAD OF SECOND-TO-LAST BIRTHS ($_2q_0$)
1	1012	0.185		
2	825	0.156	807	0.252
3	751	0.117	746	0.178
4	620	0.135	620	0.116
5	493	0.134	492	0.146
6	389	0.123	388	0.129
7	263	0.118	263	0.133
8+	422	0.109	420	0.129
TOTAL	4775	0.142	3736	0.166

Table 6 Child mortality estimated from proportions dead of last and second-to-last births by actual birthweight of current child: Bamako, Mali 1985

Actual Birthweight (grammes)	Last Births N	Last Births Proportions dead ($_2q_0$)	Second-to-last Births N	Second-to-last Births Proportions dead ($_5q_0$)
1500-1999	76	.197	54	.241
2000-2499	409	.161	299	.171
2500-2999	1389	.153	1079	.181
3000-3499	1827	.136	1407	.158
3500-3999	607	.104	513	.135
4000 or more	98	.092	86	.140
Total	4406	.139	3438	.165

TABLE 7 VALUES OF THE RATIO q (2)/q (3) FOR DIFFERENT LEVELS OF MORTALITY IN THE GENERAL STANDARD MODEL LIFE TABLE

Alpha	q (2)	q (3)	q (2) / q (3)
-1.0	0.0314	0.0352	0.8904
-0.8	0.0461	0.0516	0.8920
-0.6	0.0672	0.0751	0.8944
-0.4	0.0970	0.1081	0.8978
-0.2	0.1382	0.1531	0.9025
0.0	0.1930	0.2124	0.9087
0.2	0.2630	0.2869	0.9166
0.4	0.3474	0.3751	0.9261

Notes: Alpha is the parameter used to specify the level of mortality in Brass's relational system of model life tables. The central value of 0.0 implies an infant mortality rate of 150 per 1000

TABLE 8 POSSIBLE COMBINATIONS FOR THE SURVIVAL STATUS OF TWO SUCCESSIVE CHILDREN

YOUNGER CHILD	OLDER CHILD ALIVE	DEAD	TOTAL
ALIVE	A	B	A + B
DEAD	C	D	C + D
TOTAL	A + C	B + D	A + B + C + D

TABLE 9 THE EFFECT OF DIFFERENT DEGREES OF DEPENDENCE BETWEEN THE SURVIVAL OF SUCCESSIVE CHILDREN ON THE INDEX OF EARLY CHILDHOOD MORTALITY BASED ON THE PROPORTIONS OF PREVIOUS CHILDREN DEAD OBTAINED FROM MOTHERS WITH A SUBSEQUENT SURVIVING CHILD AGED ONE YEAR

	PERCENTAGE UNDERESTIMATION OF EARLY CHILDHOOD MORTALITY				
	ALPHA:	-0.4	-0.2	0.0	0.2
f	CORRESPONDING INFANT MORTALITY RATE PER 1000	73	106	150	208
1.0		0	0	0	0
1.5		3.96	5.91	8.82	13.16
2.0		7.93	11.83	17.65	26.33
2.5		11.89	17.74	26.47	39.49
3.0		15.86	23.66	35.29	52.65

Notes: The factor f describes the degree of dependence between the survival status of two successive children born to the same mother. When f = 1, the two outcomes are completely independent

TABLE 10 ADJUSTMENT FACTORS FOR THE CORRECTION OF THE PROPORTIONS OF PREVIOUS CHILDREN DEAD WHEN THE REPORTS ARE OBTAINED FROM MOTHERS WITH SUBSEQUENT SURVIVING CHILDREN

	ALPHA:	-0.4	-0.2	0.0	0.2
	CORRESPONDING INFANT MORTALITY RATE PER 1000	73	106	150	208
f					
1.0		1.00	1.00	1.00	1.00
1.5		1.04	1.06	1.10	1.15
2.0		1.09	1.13	1.21	1.36
2.5		1.13	1.22	1.36	1.65
3.0		1.19	1.31	1.54	2.11

Notes: In applications, the proportions of previous children dead are multiplied by the adjustment factors to obtain revised values for the index of early child mortality. Further details in the text.

TABLE 11. CHILDHOOD MORTALITY ESTIMATES FROM THE REPORTS ON TOTAL CHILDREN EVER-BORNE AND SURVIVING

AGE OF WOMAN	NO. OF WOMEN (1)	AVERAGE PARITY (2)	x (3)	$1000\,{}_{0}q_{x}$ (4)	ALPHA (5)	T (6)	ALPHA * (7)	$1000\,{}_{5}q_{0}$ (8)
15-19	1619	0.129	1	54.5	- 0.560	1.17	- 0.771	48.7
20-24	1597	0.864	2	58.8	- 0.671	2.52	- 0.753	50.4
25-29	1292	1.940	3	59.3	- 0.727	4.38		52.9
30-34	1051	3.143	5	67.9	- 0.708	6.56		54.8
35-39	822	4.395	10	84.2	- 0.643	8.56		62.0
40-44	650	5.298	15	87.6	- 0.658	11.60		60.2
45-49	506	6.271	20	11.2	- 0.580	14.84		69.8

Notes:

1. Mortality estimates derived using Brass's weights to convert the proportions dead to ${}_{n}q_{x}$ values.

2. Alpha values obtained by subtracting the logits of the l(x) values in the General Standard life table from the logits of the l(x) values obtained by the indirect estimation procedure.

3. Adjusted alpha values for the women aged 15-24 were obtained by extrapolation of the trend in alpha established using data from the older women (see Figure and text for clarification).

4. T is the time location of each estimate in years before the survey.

TABLE 12. SUMMARY OF CHILDHOOD MORTALITY ESTIMATES FROM DIFFERENT METHODS

SOURCE	x	1000 $_xq_0$		T
A) Brass indirect estimates from women 15-19 and 20-24:				
o Unadjusted	1	54.5	-0.560	1.2
	2	58.6	-0.671	2.5
o Adjusted	1	48.7	-0.771	1.2
	2	50.4	-0.753	2.5
B) Life table for all births in 5 years before the survey	1	40.1	-0.721	2.5
	2	47.8	-0.781	2.5
C) Preceding birth technique applied to reports from women with a live birth in the 24 months before the survey				
o Preceding birth (N=1056)	3.6	57.8	-0.761	2.1
	2	49.6	-0.761	2.1
o Second-to-last (N=680)	6.6	75.0	-0.673	3.6
	2	58.6	-0.673	3.6

Notes:

1. Brass's General Standard life table has been used throughout.

TABLE 13 LIFE TABLE ESTIMATES OF MORTALITY BY AGE 2 FOR CHILDREN BORN IN SPECIFIED TIME PERIODS

TIME PERIOD	$1000\,{}_2q_0$ (N OF BIRTHS)	TIME PERIOD	$1000\,{}_2q_0$ (N OF BIRTHS)
1986-7	38.6 (1503)		
1984-5	44.2 (1629)	1983 - 7	47.8 (3910)
1982-3	55.8 (1593)		
1980-1	39.2 (1479)	1978 - 82	40.6 (3518)
1978-9	35.9 (1224)		
1976-7	36.9 (1057)		
1974-5	39.0 (872)	1973 - 7	37.9 (2267)
1972-3	37.9 (686)		
1970-1	24.4 (491)		

Notes:

1. Calculated directly from the dates of birth and dates of death of the three last born children.

FIGURE 1 CHILDHOOD MORTALITY ESTIMATES OBTAINED BY TWO METHODS COMPARED
A = ADJUSTED MULTIPLIER TECHNIQUE
B = SURVIVAL OF LAST AND PENULTIMATE BIRTHS

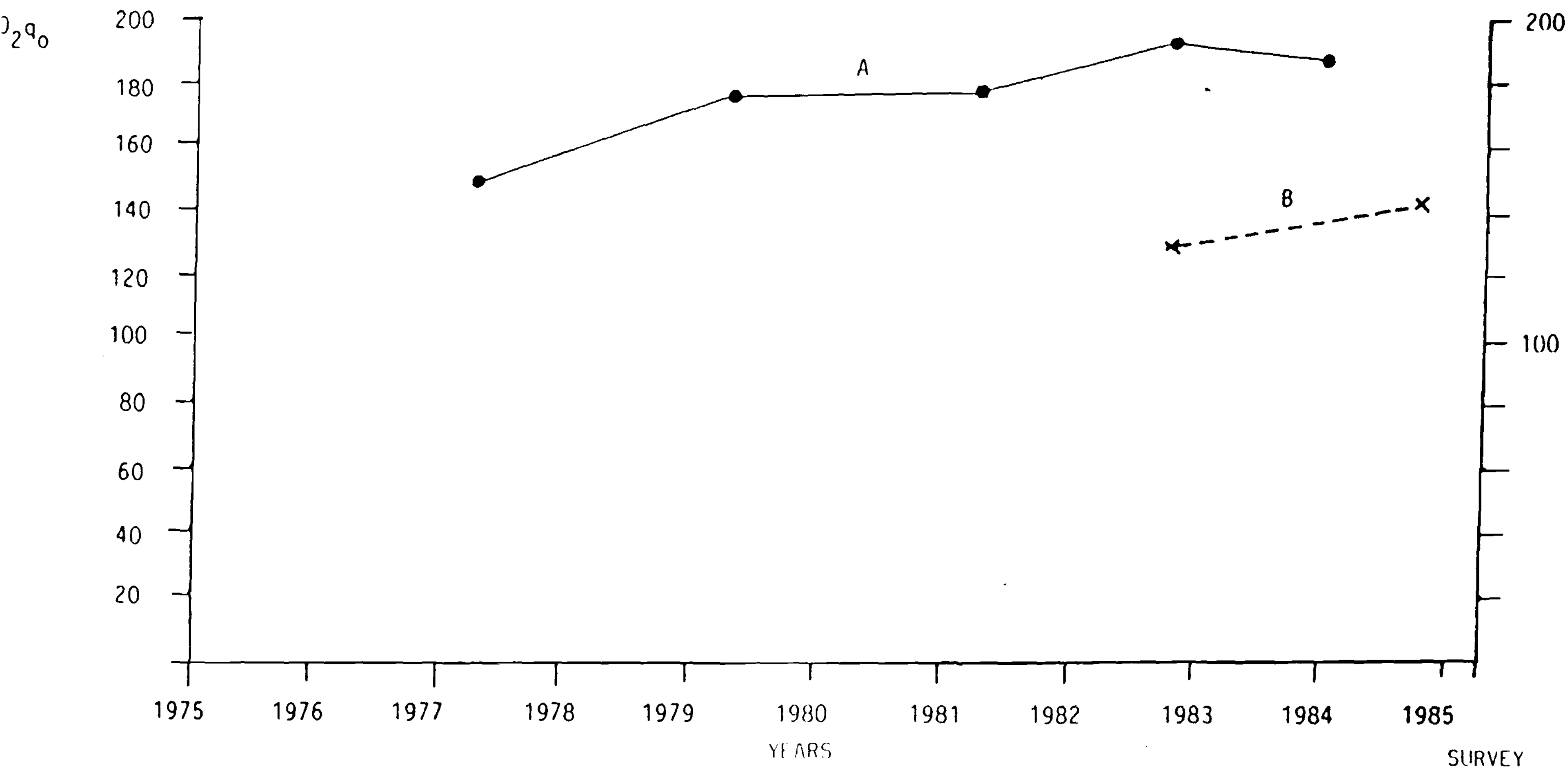


FIGURE 2

FACTORS OF DEPENDENCE BETWEEN THE PROBABILITIES OF DEATH OF SUCCESIVE SIBLINGS BY AGE OF THE YOUNGER FOR SELECTED COUNTRIES

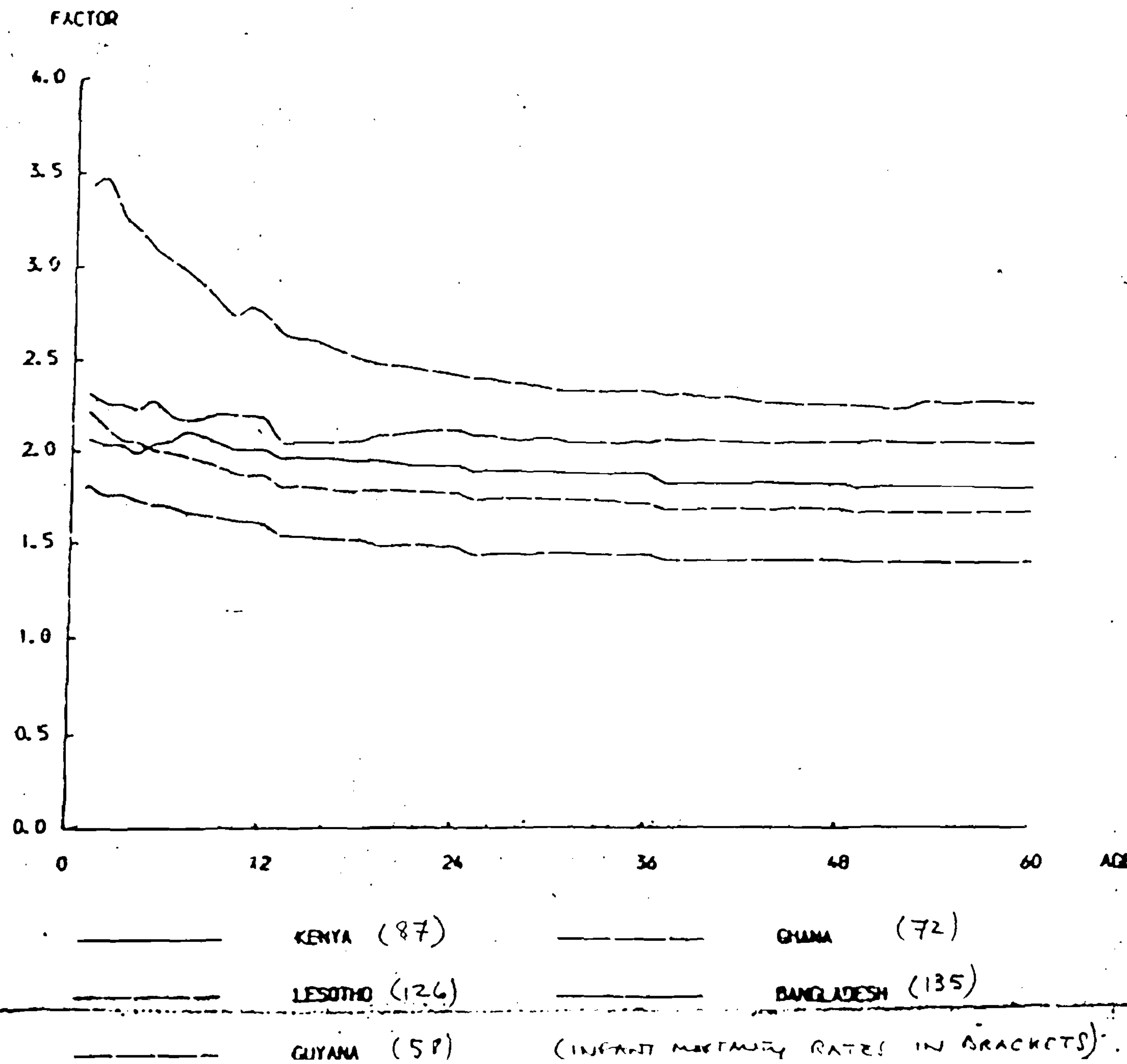
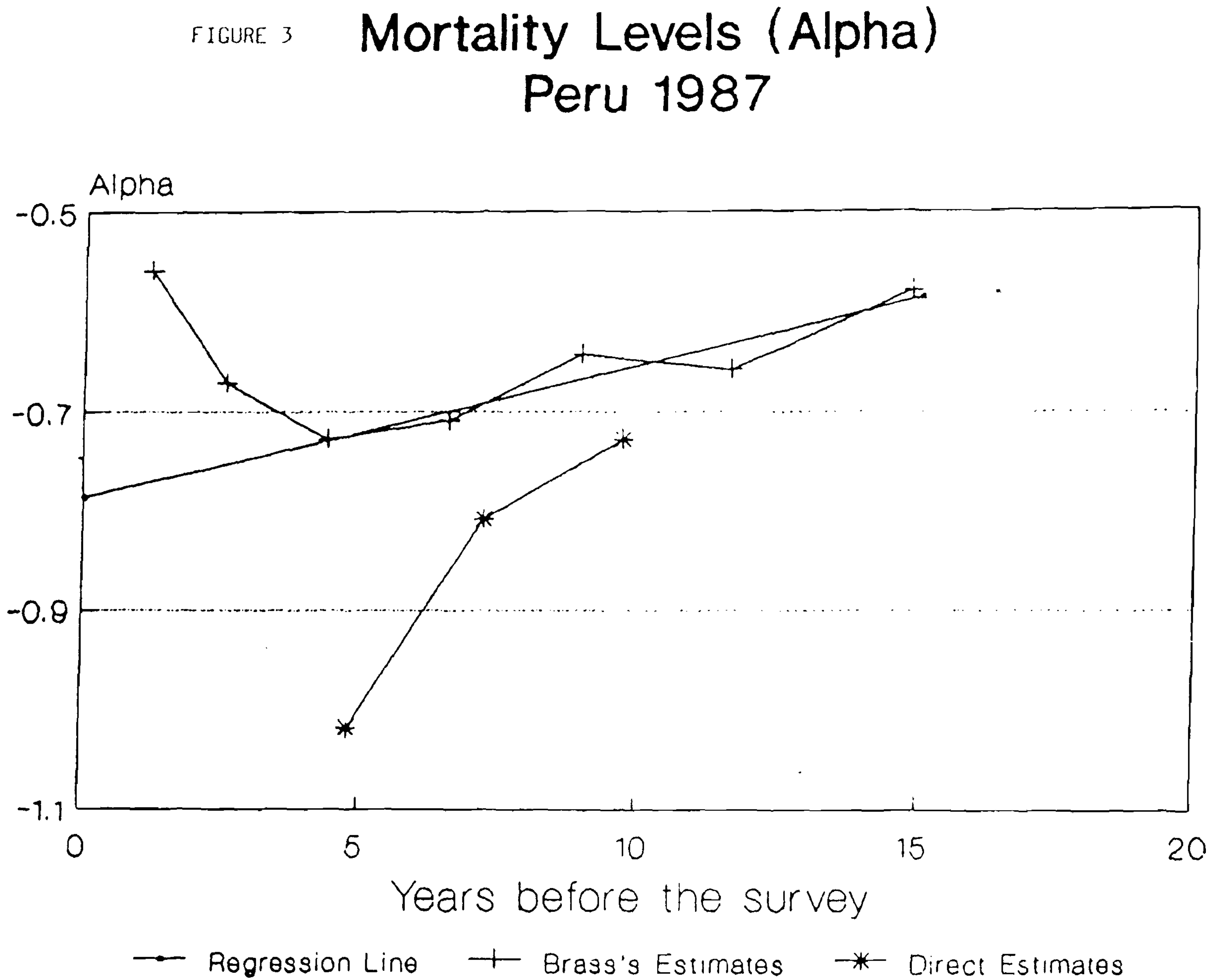


FIGURE 3


Mortality Levels (Alpha)
Peru 1987
Alpha
-0.5
-0.7
-0.9
-1.1
0
5
10
15
20
Years before the survey
Regression Line
Brass's Estimates
Direct Estimates

APPENDIX A

		MERE							GROSSESSE ACTUELLE										DECES MATERNEL		SORTIE		
No.	Date et heure d'entrée	Nom, prénom et code CPN	Age	Nbre. d'enfants nés-vivants ?	Nbre. d'enfants encore vivants ?	Nbre. d'avortements et morts-nés ?	Sexe du dernier enfant (M/F) ?	Dernier né-vivant en vie (O/N) ?	Date et heure d'accouchement ?	Lieu d'accouchement (M/D/H/A) ?	A terme ou prématuré (AT/P) ?	Accouchement (N/S/F/V/C/A) ?	Résultat (NV/MN/FC) ?	Unique ou multiple (U/M) ?	Sexe (M/F) ?	Poids de naissance (g.) ?	Taille de l'enfant (cm.) ?	BCG (O/N) ?	Période (Av/Pd/Ap) ?	Cause ?	Date et heure de sortie ?	Enfant encore en vie (O/N) ?	Observations
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)	(18)	(19)	(20)	(21)	(22)	(23)	(24)