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Ethnic Differentials in Early Childhood Mortality in Nepal

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CONTENTS

TABLES

FIGURES

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ABSTRACT

This paper investigates the effects of ethnicity on early childhood mortality in Nepal, based on data from the 1976 Nepal Fertility Survey, which was part of the World Fertility Survey. The approach is through a series of hazard models, which incorporate ethnicity, year of birth, mother's literacy, father's literacy, rural-urban residence, region, sex, maternal age, survivorship of previous birth, previous birth interval, and breastfeeding as covariates. The analysis indicates that ethnic differentials in early childhood mortality are not explained by the other covariates. An implication is that future studies of ethnic differentials need to collect more detailed information on circumstances of childbirth and childrearing practices relating to health. A byproduct of the analysis is an improved specification of breastfeeding as an age-varying covariate that indicates that breastfeeding (relative to not breastfeeding) reduces age-specific mortality risks during the first two years of life by about 75 percent.

ETHNIC DIFFERENTIALS IN EARLY CHILDHOOD MORTALITY IN NEPAL

Recent advances in multivariate analysis of childhood mortality have resulted in a rapidly improving understanding of the interplay of the socioeconomic, demographic, biomedical, and environmental determinants of child mortality. The literature on determinants of child mortality has been summarized by Mosley and Chen (1984), Preston (1985), Mensch, Lentzner, and Preston (1985), and Pebley and Millman (1986).

Briefly, some principal findings relevant to the present analysis are the following: Parental education, especially mother's education, tends to reduce child mortality. Urban residence, which tends to lower child mortality in bivariate analyses, usually has little effect once parental education and major demographic factors are controlled. Sex of child is usually important during the first year of life, with males tending to higher mortality, but usually unimportant after the first year. First births tend to have higher mortality than later births, but birth order makes little difference among later births, once maternal age, previous birth interval, and other demographic factors are controlled. Child mortality tends to be higher at the youngest and oldest maternal ages (especially the youngest ages) and lower in between. Birth intervals, both immediately preceding and immediately following the birth of the index child, have substantial effects on child mortality, with longer intervals tending to reduce mortality. Survivorship of previous birth tends to lower mortality risks for the index child. The most important single piece of research contributing to these findings is that by Hobcraft, McDonald, and Rutstein (1985), who present results of fitting proportional hazard models for 39 countries (including Nepal) covered by the World Fertility Survey.

Breastfeeding also tends to lower child mortality, and it is sometimes included in addition to birth interval covariates in hazard models. Although breastfeeding tends to lengthen birth intervals, the effect of birth interval is not completely explained away by breastfeeding; both breastfeeding and birth interval tend independently to lower child mortality (Pebley and Stupp, 1986; Palloni and Millman, 1986).

A few multivariate analyses in the demographic literature have included not only the above-mentioned socioeconomic and demographic covariates (or at least some of them) but also such biomedical covariates as birthweight, whether the child was born in a hospital, and whether the child was born subsequent to a health intervention program, and such environmental covariates as water supply and toilet facilities (see, for example, DaVanzo, Butz, and Habicht, 1983; Martin et al., 1983; Pebley and Stupp, 1986). Most of these additional covariates have been found to have substantial effects on child mortality, even after the socioeconomic and demographic variables just mentioned (or at least some of them) are controlled. This finding is not surprising, given that progress in medicine and public health is known to have been extremely important in bringing down mortality in developing countries (Davis, 1956; Preston, 1975, 1980).

A few studies have examined ethnic differentials in child mortality, controlling for covariates. DaVanzo, Butz, and Habicht (1983) found that in Malaysia substantial child mortality differences among Malays, Chinese, and Indians persisted after introduction of controls for socioeconomic, demographic, biomedical, and environmental variables. The analysis was based on both linear probability and logit regression models. Mensch, Lentzner, and Preston (1985) also found for eleven countries (including Nepal) that ethnic differentials in child mortality tended to persist after the introduction of controls. However, in this case the controls were less comprehensive (controls for demographic covariates were largely lacking), and the authors used an indirect measure of child mortality that is less than ideal for such an analysis. As Preston (1985) has noted, direct measures of child mortality tend to be superior to indirect measures in the World Fertility Surveys that have provided the basis for many recent multivariate analyses of child mortality.

The apparent failure of socioeconomic, demographic, biomedical, and environmental variables (or at least the limited number of such variables included in the above-mentioned studies) to explain ethnic differentials in child mortality is intriguing and merits further investigation. In this paper we fit hazard models to data from the 1976 Nepal Fertility Survey to investigate the magnitude and causes of ethnic differentials in child mortality in Nepal. The analysis is more detailed and refined than the earlier analysis by Mensch, Lentzner, and Preston (1985).

DATA AND METHODS

Previous studies have demonstrated that the 1976 Nepal Fertility Survey (NFS), which was part of the World Fertility Survey, is the best source of child mortality data in Nepal (Goldman, Coale, and Weinstein, 1979; Thapa and Retherford, 1982; Gubhaju et al., 1987). We have therefore restricted our investigation to this data source. The NFS included 5,940 ever married women aged 15-49. Child mortality data are contained in the birth histories. The survey also collected considerable information on demographic covariates of child mortality as well as adequate information on socioeconomic covariates. Data on relevant biomedical and environmental covariates are lacking and perforce omitted from our analysis. Details of survey methodology and sample design of the NFS are contained in the First Report of the NFS (Nepal FP/MCH Project, 1977).

The universe of births in our analysis is restricted to last and next-to-last births of order 2 or higher occurring during 1972-76 (i.e., during the 60-month period immediately preceding the survey). The restriction to last and next-to-last births allows inclusion of breastfeeding as a covariate; in the NFS, breastfeeding information was collected only for last and next-to-last births. The restriction to births of order 2 or higher allows inclusion of previous birth interval as a covariate. The restriction of the "window" of births to 1972-76 helps to reduce bias, as will be discussed later.

The dependent variable in our analysis is child mortality up to the age of two years, conceptualized sometimes as the set of monthly age-specific mortality risks before age 2 and sometimes as the probability of dying before age 2, denoted here as $q(2)$. The specification is in terms of multivariate life tables based on monthly risks of death. The life tables are truncated at 24 months. Originally we intended to restrict the analysis to infant mortality, i.e. child mortality up to the age of 12 months. The problem with truncating the life tables at 12 months is that Nepalese age data are characterized by severe age misreporting and age heaping (Goldman, Coale, and Weinstein, 1979). Specifically, many infant deaths that occurred at ages less than 12 months were rounded up to 12 months, and the extent of such rounding varies by ethnic group, as shown in Figure 1. In the figure, the Thakuri-Chhetri group shows considerable heaping on 12 months whereas Brahmans show comparatively little. In instances where the rounding is extensive, infant mortality tends to be severely underestimated (Thapa and Retherford, 1982; Gubhaju et al., 1987), yielding distorted estimates of ethnic differentials in infant mortality. Heaping also occurs at 24 months, which is also shown in the figure. But since mortality is quite low at 24 months compared with earlier ages, the bias in the estimated probability of dying before 24 months that results from heaping on 24 months is much less than the bias in the estimated probability of dying before 12 months that results from heaping on 12 months. Thus the choice of a 24-month cutoff.

There are many ethnic groups in Nepal, and it is not possible to examine each of them separately. The NFS classified women into 21 major ethnic groups and a residual category of "Other." To render the analysis more manageable, we have collapsed the original 22 categories into eight, based on marriage patterns, acceptability of interethnic marriages, race, religion, caste, and other cultural similarities. Bista's (1972) ethnographic descriptions of ethnic groups in Nepal were helpful in arriving at the reclassification. The collapsed groups are Rai-Limbu (also referred to as Kirate), Brahman, Satar (including Sunwar, Darwar, Mosar, Darai, and Tharu, as well as Satar), Musalman (Muslim), Newar, Tamang-Gurung-Magar, Thakuri-Chhetri, and Other. This breakdown is more elaborate than that used in the previously mentioned study by Mensch, Lentzner, and Preston (1985), who used six groups (Brahman, Newar, Chhetri, Magar, Tharu, and Other).

Because infant mortality has been falling fairly rapidly in Nepal, and at different rates in different ethnic groups, we have included as a control a covariate specifying year of birth. In the analysis, year of birth is treated as a continuous variable. To further reduce bias from differing trends in child mortality, changing relationships with covariates, and selectivity according to fucundity, we have restricted the "window" of births to the five years immediately preceding the survey. Next and next-to-last births account for 92.4 percent of all births during this five-year period.

Socioeconomic (including geographic) covariates included in the analysis include mother's literacy, father's literacy, rural-urban residence, and region (hill, terai, mountain). ("Father" in this context means mother's husband at the time of the survey or, if mother was not currently married at the time of the survey, her last husband.) Previous analyses of the determinants of child mortality using NFS data have indicated the importance of mother's education and region of residence (Thapa and Retherford, 1982; Gubhaju, 1984, 1986). The NFS did not include a question on income, but in the present context this may not be a serious omission, because previous studies in other countries indicate that the effect of income on child mortality is often greatly reduced when mother's and father's education are taken into account (DaVanzo, Butz, and Habicht, 1983; Cochrane, O'Hara, and Leslie, 1980; Mensch, Lentzner, and Preston, 1985). The socioeconomic covariates are represented by dummy variables in the analysis.

Seven demographic covariates are included in the analysis. The first of these is sex, represented as a dummy variable (1 if male, 0 otherwise). Although sex is presumably uncorrelated with ethnicity in the population, it is both subject to sampling variability and usually correlated with child mortality, so that introducing sex to minimize the effects of stochastic variation in the sex ratio seems warranted. Maternal age has been shown to have effects on child mortality and is included to control for differences in mean age at childbirth among ethnic groups. Maternal age in years is treated as a continuous variable in the analysis. Since the effect of maternal age tends to be somewhat U-shaped, maternal age squared is also included. Another demographic covariate is whether the preceding child survived to the time of the survey. This variable helps to control for family-level differences in previous child mortality, thereby clarifying the interpretation of the effect of previous birth interval on child mortality (Hobcraft, McDonald, and Rutstein, 1985:375). Previous birth interval has been shown in many studies to have strong effects on child mortality and may help explain ethnic differentials in child mortality because of possible ethnic differences in breastfeeding, which affects birth intervals as well as child mortality directly. Previous birth interval (in months) is treated as a continuous variable in the analysis. Because birth interval effects on child mortality are often nonlinear (Hobcraft, McDonald, and Rutstein, 1985), previous birth interval squared is also included. Breastfeeding is additionally included among the demographic covariates because of its effect on birth intervals, its strong direct effect on child mortality, and its tendency to vary from group to group within populations (Knodel, 1968). Since only mortality below the age of two years is considered, breastfeeding should capture much of the effect of following birth interval, which is excluded from the analysis. (Because our primary interest is ethnic differentials in child mortality rather than child mortality itself, and because of cost considerations, we decided to limit the number of birth interval variables, which, as we shall see, tend not to be correlated with ethnicity anyway.)

Breastfeeding, coded as a dummy variable, is included as an age-varying covariate. In the models, at each age of child, the risk of death depends on breastfeeding status one month earlier; thus the value of the breastfeeding variable may change for a given child as the child gets older. For the first month of life, breastfeeding is coded "yes," given that breastfeeding one month prior is meaningless when the index child is less than one month old. Since there are no doubt a few children who died in the first month for whom a "no" code would have been appropriate, the uniform assumption of "yes" probably produces a small downward bias in the estimated effect of breastfeeding. However, undoubtedly a sizeable fraction of the "no" cases were never breastfed only because they were too sickly at birth and died soon thereafter. In such instances, a "yes" code is not unreasonable, because the lack of breastfeeding is appropriately viewed as caused by child mortality rather than the other way around.

Our set of demographic covariates excludes birth order. In their review of World Fertility Survey findings for 39 countries, Hobcraft, McDonald, and Rutstein (1985) found that when other variables were controlled, child mortality tended to be higher for first births but to vary little by order for subsequent births. Since our analysis of the NFS data is restricted to births of order 2 or higher, and since preliminary analysis indicated that demographic covariates tend not to be correlated with ethnicity, inclusion of birth order seemed unnecessary. Cost considerations also played a role in our decision to exclude birth order.

The covariates are employed in a series of hazard models (for background on hazard models, see Cox, 1972; Kalbfleisch and Prentice, 1980:70-84; Trussell and Hammerslough, 1983; Tuma and Hannon, 1984:233-254). The computer program used is BMDP2L. This program treats the exposure variable (age) as continuous, and it accepts both continuous and dummy variables as covariates. The inclusion of age-varying covariates is handled very simply by specifying the covariates as functions of age. Partial likelihood is used for estimation. It has been shown that partial likelihood can be treated like the usual maximum likelihood (Tuma and Hannan, 1984:244-247).

Some covariates of infant and child mortality are known to have different effects on mortality at different ages. For example, environmental factors are known to affect mortality during late infancy and subsequent childhood more than during early infancy. More general forms of hazard models can be used to estimate effects that change with age, but application of such models requires more accurate age reporting than is found in the Nepal data. An estimated effect in this report may be considered as an average effect in the age interval from birth to two years.

Based on findings from previous studies mentioned earlier, we expected considerable variability in child mortality among ethnic groups in Nepal. We hypothesized that some of this variability would be explained by the socioeconomic covariates, especially mother's education, and by the demographic covariates, especially breastfeeding and previous birth interval. However, we also hypothesized that some unexplained variability would remain, just as in previous studies. As the findings in the next section show, not all of these expectations and hypotheses were borne out by the analysis.

RESULTS

Before proceeding to the hazard analysis, we computed simple life tables for each of the eight ethnic groups, yielding estimates of q(2), the probability of dying before two years of age. (Although the argument of q(2) is denoted in years, the life tables themselves were computed from monthly risks of death.) Summary statistics from these life tables are shown in Table 1. The table shows considerable ethnic variation, with q(2) varying from 32 for Rai-Limbu and 99 for Brahman to 133 for Tamang-Gurung-Magar. Standard errors of these values are, however, fairly large, so that most of the ethnic differences in q(2) are not statistically significant. In this regard it should be noted that the estimated q(2) for Rai-Limbu, an outlier that is exceptionally low, is based on very few deaths and must be considered unreliable. Note that the estimates of q(2) in this table are considerably lower than they would be if first order births, which have higher mortality, were included in the analysis.

Whereas Table 1 presents summary values of q(2), Figure 1 is more detailed and presents for each ethnic group the life table survivorship function at intervals of one month. The status of Rai-Limbu as an outlier is amply evident; because of the small number of cases and consequent high sampling variability, there are no deaths at all between 8 months and 2 years of age. The other ethnic groups show more reasonable patterns, with cumulative survivorship declining regularly with age.

Table 2 shows the results of fitting five hazard models. Model I incorporates only ethnicity and year of birth as covariates. Model II is the full model, including not only ethnicity and year of birth but also socioeconomic and demographic covariates. Relative to Model II, Model III deletes the ethnicity covariates, Model IV deletes the socioeconomic covariates, and Model V deletes the demographic covariates. Comparison of the five models yields assessments of how the effects of ethnicity on child mortality are affected by each block of covariates and by all covariates simultaneously. Table 3, to which we shall return later, contains an alternative, more compact presentation of results from the models, and Tables 4-7 show ethnic differences in each covariate that are helpful in interpreting findings from the models.

Before proceeding to these findings, a few comments on the format of Table 2 are in order. The effects of the various covariates are expressed as relative risks. For example, in the full model (Model II), the relative risk of 1.3611 for Thakuri-Chhetri means that the risk of dying in any given one month age interval, controlling for other covariates in the model, is 1.3611 times higher for Thakuri-Chhetri than for Brahmans, who are the reference category. The relative risk of 0.9292 for year of birth means that each one-year increment in year of birth reduces age-specific mortality risks by a factor of .9292. The relative risk of 1.6143 for "mother illiterate" means that age-specific mortality risks are 1.6143 times higher for children of illiterate mothers than for children of literate mothers, who are the reference category.

The relative risks are calculated as $\exp(B_i)$, where B_i is the underlying parameter estimate (not shown) for the ith covariate. The symbols indicating level of significance (plus, asterisk, or pound marks) refer to the underlying parameter estimates. In the case of continuous variables and dichotomous discrete variables, the meaning of the significance levels is clear enough (see footnote to Table 2). However, for discrete variables with three or more categories, the significance tests should ideally include all pairwise comparisons rather than simply comparisons of each category with the reference category. Pairwise comparisons are shown in the case of region. However, in the case of ethnicity, with eight categories, the number of pairwise comparisons is excessive. Our strategy was instead to choose the reference category as the ethnic group with the lowest child mortality, which we took as Brahmans since the estimate for Rai-Limbu was considered unreliable (see Table 3). Thus, for a given ethnic group, a parameter estimate significantly different from zero means that the ethnic group in question differs significantly from Brahmans. Choosing the group with the lowest child mortality as the reference category maximizes the number of significant differences. The observed levels of significance for ethnicity in Table 2 are low considering the number of cases, because the data suffer from large variances resulting from considerable statistical noise generated by age misreporting. Thus it could be argued that the estimated effects, interpreted as average effects over a two-year age interval, are more statistically significant than they appear to be at first sight from the standard tests.

Proceeding now to the findings in Table 2, we find from Model I that ethnicity, with only year of birth controlled, has a considerable impact on age-specific mortality risks. Not surprisingly, the picture is rather similar to that of Table 1, based on simple life tables. There are, however, slight differences in the order of ethnic groups by mortality level (the relative positions of Satar and Musalman are reversed), owing to the control for year of birth as well as the simplifying assumptions (proportional hazards) embodied in Model I.

Model II is the complete model. Surprisingly, the controls for socioeconomic and demographic covariates make very little difference in the relative risks by ethnicity. Comparison of Model II with Model I shows that relative risks increase slightly for four ethnic groups and decrease slightly for three. The relative risk for year of birth changes hardly at all. The overall picture is that the effects of ethnicity on child mortality are for the most part statistically independent of the socioeconomic and demographic covariates included in Model II. In other words, the socioeconomic and demographic covariates included in the model are of little help in explaining ethnic differentials in child mortality.

This is not to say, of course, that the socioeconomic and demographic covariates do not affect child mortality itself. Model II shows very clearly that these variables have considerable effects on child mortality, which we review briefly now.

Consistent with previous studies, Model II shows that mother's illiteracy, with a relative risk of 1.61, substantially increases child mortality, but father's illiteracy, with a relative risk of 1.02, makes virtually no difference.

Geographic region has less impact than literacy. Interestingly, with other factors controlled, the terai region has, by a slight margin, the lowest child mortality of the three regions, whereas earlier studies, without controls or with only a few controls, have shown that the hill region has the lowest mortality (see, for example, Thapa and Retherford, 1982; Gubhaju, 1986). The finding that the terai has the lowest mortality when other factors are controlled is not implausible, since the terai has the best health services network among the three regions. In Model II, the mountain region has the highest child mortality, consistent with earlier studies. It must be born in mind, of course, that Model II excludes first births and therefore is not strictly comparable with earlier studies. Moreover, the regional differences in Model II are not statistically significant.

Another surprise is that rural residence, relative to urban, tends to result in lower mortality. The relative risk for rural, with other variables controlled, is .82. Our expectation was that rural residence would make no difference, given that previous studies have shown that when mother's education is controlled, rural-urban residence usually makes little difference in child mortality (Hobcraft, McDonald, and Rutstein, 1984). To the extent that residence does make a difference, we expected that rural residence would increase relative risk, as Mensch, Lentzner, and Preston (1985) found, since health services tend to be concentrated in urban areas. On the other hand, public sanitation tended still to be very poor in Nepali cities in 1976, when the NFS was taken. The situation may have been somewhat like that which typically obtained in 19th century Europe, where urban mortality was considerably higher than rural mortality (Davis, 1965). Mensch, Lentzner, and Preston (1985):255) also note that an underlying urban hazard sometimes appears in contemporary developing countries when other factors are controlled. Again, however, caution is required in interpreting this finding in the case of Nepal, since the rural-urban difference in Model II is not statistically significant, reflecting the very small size of the urban component of the NFS sample.

We turn next to the demographic covariates. Sex of child has virtually no effect on child mortality. Maternal age and maternal age squared have effects in the expected direction. The effect of maternal age is substantial (a 7 percent reduction in child mortality for each additional year of maternal age), but the underlying parameter estimate is not statistically significant. This finding is fairly consistent with that of Hobcraft, McDonald, and Rutstein (1985), who found that when demographic covariates, particularly birthspacing covariates, were controlled, the effects of maternal age were largely removed. Note that the relative risk for maternal age squared, though also not statistically significant, is larger than it looks, since the units in this case are squared age units. Survivorship of previous birth, birth interval, birth interval squared, and breastfeeding also have effects in the expected direction. The effects are mostly large, and the parameter estimates tend to be highly significant.

The huge effect of breastfeeding is especially noteworthy; the child mortality risk for breastfed children is only 24 percent as high as the child mortality risk for children who are not breastfed. (This relative risk is further reduced to 13 percent when breastfeeding is coded alternatively as "no" for children aged less than one month; however, the estimated ethnic differentials in child mortality are affected hardly at all by the alternative choice of breastfeeding specification.) This result suggests that breastfeeding, which is strongly correlated with birth interval, probably captures much of the effect of following birth interval, which was not included in our model.

Some caution is nevertheless necessary in interpreting these findings for breastfeeding. The information on length of breastfeeding was collected in a different part of the questionnaire than the birth history, and the lengths of breastfeeding are severely heaped on 6, 12, and 18 months (Goldman, Coale and Weinstein, 1979). Some of the apparent effect of breastfeeding may result from inconsistent heaping of ages and lengths of breastfeeding. For example, there may have been a tendency to exaggerate reported length of breastfeeding for surviving children but not for children who died. On the other hand, the finding of a huge breastfeeding effect is plausible, given the unsanitary and poorly nutritious alternatives to breastfeeding that generally prevail in Nepal. Further research is needed on this question.

Another indication of the power of the socioeconomic and demographic variables in explaining variation in child mortality, if not ethnic differentials in child mortality, is found in the last panel of the table on likelihood statistics, which yield a global test of the difference between Model I and Model II. The difference between these two models is highly significant.

Model III provides another way of looking at the effect of ethnicity on child mortality, by deleting the ethnicity covariates from Model II. Except for the effects of region, which become slightly stronger, relative risks for the remaining coefficients are changed hardly at all by this deletion. This comparison between Models II and III indicates that the effects of the socioeconomic and demographic covariates on child mortality do not depend at all on ethnicity. Moreover, the likelihood statistics indicate that the difference in global fit between Models II and III is not statistically significant. By this measure, ethnicity adds little to the explanation of child mortality, once other covariates are controlled, despite some statistically significant effects of ethnicity on child mortality in Model II. We interpret this to mean that although one ethnic group (Tamang-Gurung-Magar) differs significantly from Brahmans and another group (Thakuri-Chhetri) differs marginally significantly from Brahmans in Model II, these two groups are not proportionately numerous enough for ethnic differences in child mortality to contribute much to overall variability in child mortality, once other covariates are controlled.

Model IV deletes socioeconomic covariates from Model II. Relative risks for the various ethnic groups tend to increase slightly, indicating that socioeconomic factors explain a small amount of ethnic variation in child mortality, once year of birth and demographic covariates are controlled. This occurs primarily because the key variable, mother's literacy, which tends to lower child mortality, is positively associated with being Brahman (see Table 4). In Model IV, the relative risks for year of birth and the demographic covariates are affected virtually not at all by the deletion of the socioeconomic covariates, indicating that the effects of demographic covariates on child mortality do not depend on the socioeconomic covariates. The likelihood statistics indicate that the difference in global fit between Models II and IV is not statistically significant. Thus, like ethnicity, the socioeconomic covariates add little to the explanation of variation in child mortality, once other covariates are controlled. We interpret this to mean that although literacy effects on child mortality are large and statistically significant in Model II, literate mothers are not proportionately numerous enough for illiterate-literate differences in child mortality to contribute much to overall variability in child mortality, once other covariates are controlled.

Model V deletes demographic covariates from Model II. When this is done, relative risks for the ethnic groups tend to decrease slightly. Table 6 suggests that this occurs because Brahmans, despite their comparatively low child mortality, tend to have somewhat shorter birth intervals than the other ethnic groups. Because shorter birth intervals tend to cause higher child mortality, child mortality differences between Brahmans and the other ethnic groups increase slightly when birth intervals are equalized through statistical controls. Interestingly, breastfeeding varies little by ethnic group (Table 7) and therefore does little to explain ethnic differentials in child mortality, contrary to expectation. Although the demographic variables influence relative risks by ethnicity very little, they make a great deal of difference in model fit. The likelihood statistics show that the difference in global fit between Models II and V is highly significant. It is clear that, compared with the ethnicity and socioeconomic covariates, the demographic covariates account for the lion's share of variability in child mortality in Models II, III, and IV.

The failure of the socioeconomic and demographic covariates to explain ethnic differentials in child mortality is shown also by an alternative presentation of the results in Table 3. In each model, estimates of q(2) by ethnicity are obtained by setting all covariates except ethnicity at Brahman values. These estimates are seen to be remarkably stable regardless of which sets of controls are incorporated in the model. As already mentioned, only the deletion of mother's illiteracy, when the block of socioeconomic covariates is excluded in Model IV, seems to make a noticeable difference in the estimates, and it makes a small one at that. In Model IV, those ethnic groups with low literacy of mother (Rai-Limbu, Satar, Musalman, Tamang-Gurung-Magar, Thakuri-Chhetri, and "Other" in Table 4) tend to show higher estimates of child mortality, relative to corresponding estimates from the other models that control for mother's illiteracy. Thus mother's illiteracy explains some of the ethnic differentials (relative to Brahmans) in child mortality.

Tables 4-7 show values for each of the principal covariates by ethnicity. The ethnic pattern of mother's literacy in Table 4 has already been touched upon. Table 4 also shows that Newars stand out as more urbanized than the other ethnic groups, and that the ethnic groups vary considerably in their distribution among the three regions. Table 5 shows sex ratios at birth by ethnicity. There is considerable stochastic variation in the sex ratio, justifying the inclusion of sex among the demographic covariates. Table 6 has already been touched upon in connection with the interpretation of the estimated effects of demographic covariates on ethnic differentials in child mortality. Table 7 shows that ethnic differences in breastfeeding are small and not statistically significant. Not shown in Tables 4-7 is mean maternal age at childbirth by ethnicity for the universe of births considered in the models. The mean maternal ages are 30.1 for Rai-Limbu, 27.3 for Satar, 28.6 for Newar, 27.2 for Brahman, 28.0 for Thakuri-Chhetri, 29.2 for Tamang-Gurung-Magar, 27.5 for Musalman, and 27.9 for "Other." These differences, though not negligible, do not appear to be very important in explaining ethnic differentials in child mortality.

CONCLUSION

Beyond a slight effect of mother's literacy, none of the socioeconomic or demographic covariates included in this analysis contributes appreciably to explaining ethnic differentials in child mortality in Nepal. Thus, despite better measurement of variables, more variables, and an improved model specification, we reach basically the same conclusion as Mensch, Lentzner, and Preston (1985), who also found that ethnic differentials in child mortality persist in Nepal when other variables are controlled.

Mensch and her colleagues did not include breastfeeding or previous birth interval in their models, and we expected that these covariates would explain ethnic differentials in child mortality to some extent. This expectation was based on Knodel's (1968) earlier work using 19th century German data, which indicated that variations in length of breastfeeding and hence birth interval helped considerably to explain differences in child mortality among the villages in his study. In Nepal, however, breastfeeding varies little by ethnicity, at least below the age of two years. It therefore contributes little to the sought-after explanation of ethnic differentials in child mortality, despite very large effects on child mortality itself. In this regard, an interesting aspect of our hazard models is an improved specification of breastfeeding as an age-varying covariate, yielding an estimated effect of breastfeeding amounting to a 76 percent reduction of age-specific mortality risk.

A fuller understanding of ethnic differentials in child mortality will require that future surveys collect considerably more information on circumstances of childbirth and on childrearing practices relating to health than did the NFS. More specifically, it would be a useful next step to collect detailed information on such variables as attendance at birth (e.g., hospital or home delivery), access to and utilization of maternal and child health services, treatment of umbilical cord, birthweight, nutrition, food preparation practices, toilet facilities, personal hygiene, water supply, vaccinations, and treatment of childhood illnesses, particularly infant diarrhea. Some previous studies have shown that ethnic differentials in child mortality persist even when some of these variables have been controlled, but no study has come close to controlling adequately for all of them. A final point involves sample size. Particularly in countries with lower rates of child death than Nepal, future surveys investigating child mortality should include samples larger than those typically found in the World Fertility Survey program, in order to minimize ambiguities in estimated effects arising from sampling variability.

Table 1: Child mortality by ethnic group

Ethnic group	q(2)	Births
Rai-Limbu	32 (16)	143
Brahman	99 (18)	398
Satar	91 (16)	450
Musalman	106 (29)	150
Newar	98 (24)	191
Tamang-Gurung-Magar	133 (14)	769
Thakuri-Chhetri	132 (13)	836
Other	108 (10)	1140

Notes: Child mortality is measured as the probability of dying before the age of two years, q(2), per thousand live births. q(2) is estimated by the life table method (Smith, 1980), based on monthly risks of death adjusted for censoring. Standard errors of the estimates of q(2), calculated by Greenwood's formula (Kalbfleisch and Prentice, 1980:10-16), are shown in parentheses. In this paper we have used life table computer programs from the BMDP statistical programs package.

In this table and throughout the remainder of this paper, unless otherwise specified, the universe of births is last and next-to-last births of order 2 or higher occurring during 1972-76 (i.e., during the 60-month period immediately preceding the survey).

Table 2: Effects of ethnicity on child mortality: hazard model estimates of relative risk

Covariate	Model I	Model II	Model III	Model IV	Model V
Ethnic group					
Rai-Limbu	0.6816	0.7509		0.7693	0.6779
Brahman	1.0000	1.0000		1.0000	1.0000
Satar	1.1693	1.1442		1.1905	1.1090
Musalman	1.0787	1.0955		1.1013	1.0284
Newar	1.3110	1.3409		1.3404	1.3126
Tamang-Gurung-Magar	1.3923+	1.4074*		1.4894*	1.3426+
Thakuri-Chhetri	1.4058*	1.3611+		1.4405*	1.3437+
Other	1.3122+	1.2713		1.3122+	1.2505
Year of birth	0.9210**	0.9292*	0.9316*	0.9274*	0.9212**
Socioeconomic covariates					
Mother illiterate		1.6143*	1.6415*		1.4765+
Father illiterate		1.0246	1.0463		0.9952
Rural		0.8192	0.8179		0.8033
Region					
Hill		1.0000	1.0000		1.0000
Terai		0.9697	0.9522#		1.0437
Mountain		1.1865	1.2538#		1.1654
Demographic covariates					
Sex (male)		0.9845	0.9880	0.9816	
Maternal age (years)		0.9311	0.9343	0.9373	
Maternal age squared		1.0011	1.0011	1.0010	
Previous birth surviving		0.7508**	0.7403***	0.7411***	
Prev. birth interval (months)		0.9739***	0.9743***	0.9743***	
Prev. birth interval squared		1.0001*	1.0001*	1.0001*	
Breastfeeding		0.2424***	0.2406***	0.2462***	
Likelihood statistics					
-2(log likelihood)	10223.9	10044.9	10054.5	10052.4	10219.1
Difference from II	179.0***	---	9.6	7.5	174.2***
Degrees of freedom	12	---	7	5	7

Notes: +, *, **, or *** indicates that the underlying parameter estimate (B-value) differs significantly from zero at the 10, 5, 1, or 0.1 percent level, respectively, using a two-tailed test. # indicates that the underlying parameter estimates for the terai and mountain regions differ significantly from each other at the 10 percent level using a two-tailed test. Relative risk is calculated as exp(B). The underlying life tables for the hazard models are truncated at 24 months of age. Reference groups (indicated by a row of ones) are included explicitly only for those categorical variables with more than two categories (ethnicity and region), for which the definition of the reference category would not otherwise be obvious.

Table 3: Adjusted estimates of q(2)

Ethnic group	Model I	Model II	Model IV	Model V
Rai-Limbu	65	69	73	65
Brahman	94	94	94	94
Satar	109	104	111	104
Musalman	101	100	103	97
Newar	122	121	124	122
Tamang-Gurung-Magar	129+	126*	137*	124+
Thakuri-Chhetri	130*	122+	133*	125+
Other	122+	115	122+	116

Notes: In each model, the estimates of q(2) are adjusted by setting all covariates except ethnicity at Brahman values. +, * denotes that the underlying parameter estimate (B-value) for ethnicity differs significantly from zero (the value for Brahman) at the 10 or 5 percent level, respectively.

Table 4: Percentage with specified socioeconomic or geographic characteristic, by ethnic group

Ethnic group	Mother literate	Father literate	Urban	Hill	Terai	Mountain
Rai-Limbu	7.2	64.4	1.0	93.5	6.5	0.0
Brahman	14.1	72.6	3.6	72.0	22.5	5.5
Satar	1.3	26.1	0.7	23.9	74.8	1.3
Musalman	6.6	38.0	3.2	2.2	97.8	0.0
Newar	15.1	61.8	14.9	85.4	9.8	4.8
Tamang-Gurung-Magar	2.1	37.7	1.4	76.4	13.7	9.9
Thakuri-Chhetri	4.6	51.6	0.9	54.4	25.9	19.7
Other	3.5	33.2	2.1	29.0	68.8	2.2

Note: Standard errors were not calculated for this table.

Table 5: Sex ratios at birth by ethnic group and time period

Ethnic group	1962-66	1967-71	1972-76	All
Rai-Limbu	0.85	1.12	1.13	1.06
Brahman	0.91	1.04	1.12	1.02
Satar	0.86	0.94	1.14	1.05
Musalman	1.35	0.81	1.25	1.16
Newar	1.01	0.81	1.03	0.99
Tamang-Gurung-Magar	1.05	0.99	0.98	1.03
Thakuri-Chhetri	1.01	1.10	1.06	1.03
Other	1.15	1.00	1.04	1.07

Note: "All" refers to 1962-76. Standard errors were not calculated for this table. Note, however, that variation in the sex ratio is especially great for Rai-Limbu, Musalman, and Newar, the groups with comparatively few cases as shown in Table 1. Evidently sampling variability accounts for most of the instability of sex ratios in this table.

Table 6: Percentage of previous births followed by a subsequent index birth within 24 months, by ethnic group

Ethnic group	Unadjusted	Adjusted
Rai-Limbu	23.2	28.6
Brahman	29.8	29.8
Satar	26.7	31.0*
Musalman	28.6	26.4*#
Newar	25.1	31.6#
Tamang-Gurung-Magar	23.4	30.0
Thakuri-Chhetri	28.7	29.5
Other	28.3	29.5

Notes: Unadjusted percentages were calculated by the life table method, without controls, based on monthly risks of birth. No pairwise differences between unadjusted percentages are significantly different at the 10 percent level.

Adjusted percentages, calculated by means of a proportional hazard model, are adjusted for year of birth, mother's literacy, father's literacy, rural-urban residence, region, sex, and maternal age by setting these covariates at Brahman values. As indicated by asterisks and pound marks, pairwise differences differ at the 5 percent level between Musalman and Satar and at the 10 percent level between Musalman and Newar. In this case the significance tests pertain to pairwise differences between underlying ethnicity coefficients in the hazard model.

Table 7: Percentage of births who breastfed for specified durations, by ethnic group

Ethnic group	At least 9 months		At least 15 months	
	Unadjusted	Adjusted	Unadjusted	Adjusted
Rai-Limbu	79.3	82.5	48.1	46.7
Brahman	84.4	84.4	51.1	51.1
Satar	83.4	84.4	55.8	51.2
Musalman	82.0	85.2	46.1	53.1
Newar	84.0	84.1	50.6	50.5
Tamang-Gurung-Magar	83.6	82.7	49.6	47.1
Thakuri-Chhetri	77.3	82.7	47.4	47.1
Other	81.3	83.3	48.9	48.5

Notes: Unadjusted percentages were calculated by the life table method, based on monthly risks of discontinuing breastfeeding. In computing the life tables, infant deaths and subsequent births, as well as the event of reaching the survey date, were treated as censoring events.

Adjusted percentages, calculated by means of a proportional hazard model, are adjusted for year of birth, mother's literacy, father's literacy, urban-rural residence, region, sex, and maternal age by setting these covariates at Brahman values.

No pairwise comparisons are significantly different at the 10 percent level in either unadjusted or adjusted columns. In the case of adjusted values, significance tests pertain to pairwise differences between underlying ethnicity coefficients in the hazard model; for a given difference between two ethnic groups, the test is the same whether 9 or 15 months are considered, since the test pertains to the entire life table.

FIGURE 1: REPORTED AGES AT EARLY CHILDHOOD DEATH: BRAHMAN AND THAKURI-CHHETRI

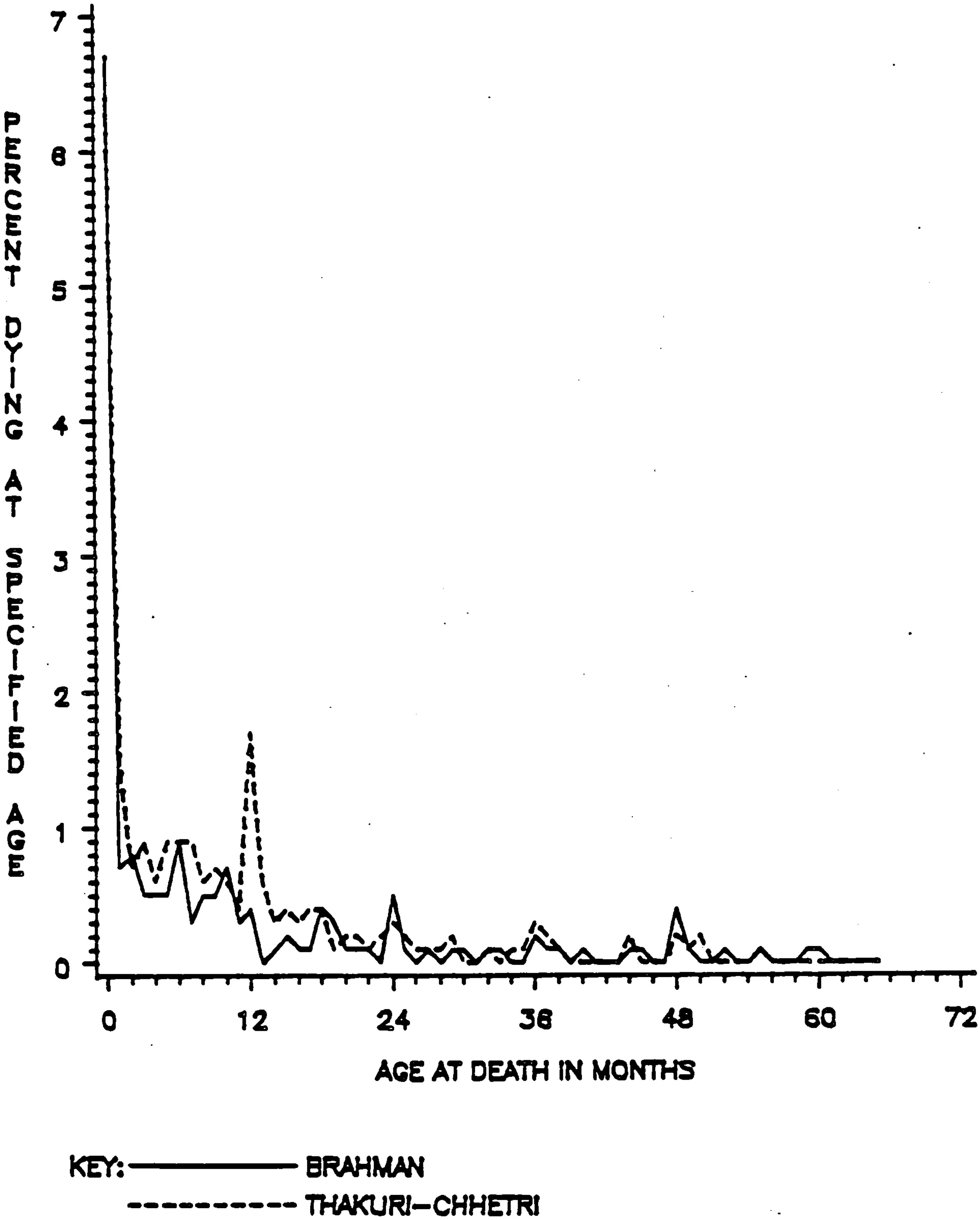
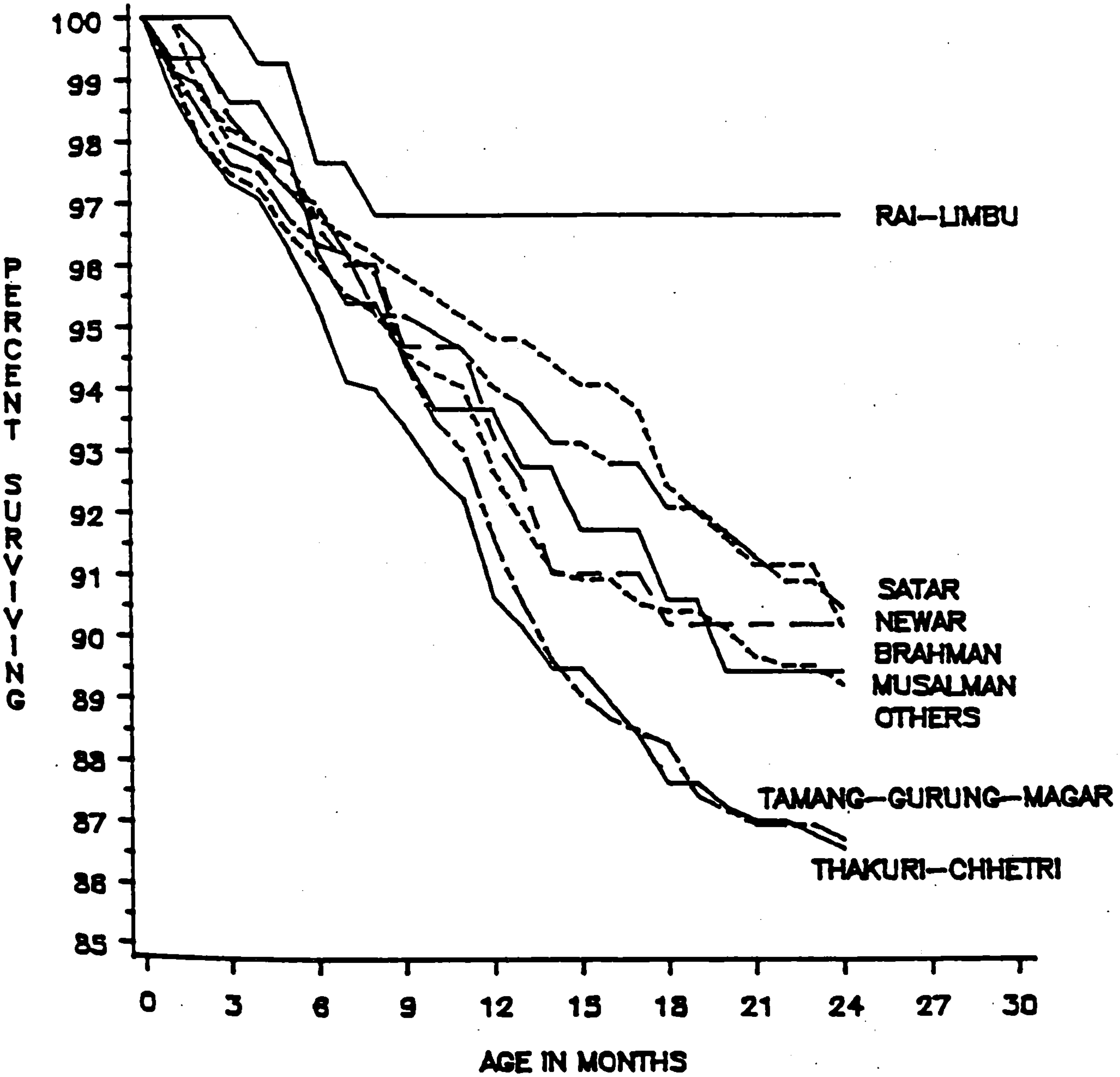


FIGURE 2: LIFE TABLE SURVIVORSHIP BY ETHNIC GROUP



NOTE: CALCULATED FROM SIMPLE LIFE TABLES WITHOUT ADJUSTMENTS FOR COVARIATES

REFERENCES

Bista, D.B. 1972. People of Nepal. Kathmandu: Ratna Pustak Bhandar.

Cochrane, S.H., D.J. O'Hara, and J. Leslie. 1980. The Effects of Education on Health. Staff Working Paper No. 405. Washington, D.C.: World Bank.

Cox, D.R. 1972. Regression models and life tables. Journal of the Royal Statistical Society B 34: 187-202.

DaVanzo, J., W.P. Butz, and J.P. Habicht. 1983. How biological and behavioural influences on moratality in Malaysia vary during the first year of life. Population Studies 37: 381-402.

DaVanzo, J., and J.P. Habicht. 1986. Infant mortality decline in Malaysia, 1946-75: The roles of changes in variables and changes in the structure of relationships. Demography 23: 143-160.

Davis, K. 1956. The amazing decline of mortality in underdeveloped areas. Journal of the American Economic Association 46: 305-318.

Davis, K. 1965. The urbanization of the human population. Scientific American 213(3): 41-53.

Goldman, N., A.J. Coale, and M. Weinstein. 1979. The Quality of Data in the Nepal Fertility Survey. Scientific Reports, No. 6. London: World Fertility Survey.

Gubhaju, B.B. 1984. Demographic and Social Correlates of Infant and Child Mortality in Nepal. Unpublished Ph.D. dissertation, Department of Demography, Australian National University, Canberra.

Gubhaju, B.B. 1986. Effect of birth spacing on infant and child mortality in rural Nepal. Journal of Biosocial Science 18: 435-447.

Gubhaju, B.B., M.K. Choe, R.D. Retherford, and S. Thapa. 1987. Infant mortality trends and differential in Nepal. Studies in Family Planning 18: 22-31.

Hobcraft, J., J.W. McDonald, and S.O. Rutstein. 1985. Demographic determinants of infant and early child mortality: a comparative analysis. Population Studies 39: 363-386.

Kalbfleisch, J.D., and R.L. Prentice. 1980. The Statistical Analysis of Failure Time Data. New York: Wiley.

Knodel, J. 1968. Infant mortality and fertility in three Bavarian villages: an analysis of family histories from the 19th century. Population Studies 22: 297-318.

Martin, L.G., J. Trussell, F.R. Salvail, and N.M. Shah. 1983. Covariates of child mortality in the Philippines, Indonesia and Pakistan: an analysis based on hazard models. Population Studies 37: 417-432.

Mensch, B., H. Lentzner, and S.H. Preston. 1985. Socio-Economic Differentials in Child Mortality in Developing Countries. New York: United Nations, Department of International Economic and Social Affairs.

Mosley, W.H., and L.C. Chen. 1984. Child Survival: Strategies for Research. Population and Development Review, Supplement to Volume 10.

Nepal FP/MCH Project. 1977. Nepal Fertility Survey 1976: First Report. Kathmandu: Ministry of Health.

Palloni, A., and S. Millman. 1986. Effects of inter-birth intervals and breastfeeding on infant and early childhood. Population Studies 40: 215-236.

Pebley, A.R., and S. Millman. 1986. Birthspacing and child survival. International Family Planning Perspectives 12: 71-79.

Pebley, A.R., and P.W. Stupp. 1986. Reproductive patterns and child mortality in Guatamala. Demography 24: 43-60.

Preston, S.H. 1975. The changing relation between mortality and level of economic development. Population Studies 29: 231-248.

Preston, S.H. 1980. Causes and consequences of mortality declines in less developed countries. In R.A. Easterlin (ed.), Population and Economic Change in Developing Countries. Chicago: University of Chicago Press.

Preston, S.H. 1985. Mortality in childhood: lessons from the WFS. In J. Cleland and J. Hobcraft (eds.), Reproductive Change in Developing Countries: Insights from the World Fertility Survey. Oxford: Oxford University Press.

Smith, D.P. 1980. Life Table Analysis. Technical Bulletin, No. 6. London: World Fertility Survey.

Thapa, S., and R.D. Retherford. 1982. Infant mortality estimates based on the 1976 Nepal Fertility Survey. Population Studies 36: 61-80.

Tuladhar, J.M. 1984. Persistance of High Fertility in Nepal. Unpublished Ph.D. dissertation. Department of Demography, Australian National University, Canberra.

Tuma, N.B., and M.T. Hannan. 1984. Social Dynamics. Orlando: Academic Press.

Trussell, J., and C. Hammerslough. 1983. A hazards-model analysis of the covariates of infant and child mortality in Sri Lanka. Demography 20: 1-26

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