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John Knodel and Apichat Chamratrithirong

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JOHN KNODEL is Research Associate at the Population Studies Center of the University of Michigan and Associate Professor of Sociology, University of Michigan. APICHAT CHAMRAT-RITHIRONG is a Research Associate at the Institute for Population and Social Research, Mahidol University.

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PREFACE

A number of persons were extremely helpful to us in the preparation of this study. The National Statistical Office of Thailand kindly provided us with unpublished data from the 1970 census and the new Survey of Population Change, which served as the basis of this report. Anuri Wanglee and Chintana Pejaranonda of that organization were helpful in answering numerous questions concerning the data. John Fulton of Brown University kindly gave us permission to cite portions of his unpublished work on Thai mortality. Ken Hill of the Committee on Demography and Population of the National Academy of Sciences and James Trussell of Princeton University provided helpful technical advice. Napaporn Burapathana and several staff members of the Institute of Population Studies, Chulalongkorn University, and the Institute for Population and Social Research, Mahidol University, provided assistance in the calculations.

ABSTRACT The application of techniques designed to estimate indirectly the levels of infant and child mortality from survey and census data on the proportion dead among children ever born permits us to expand substantially our knowledge of trends and differentials in infant and child mortality in Thailand. The registration data are seriously deficient. Although the indirect estimation techniques may underestimate mortality to some extent, comparisons with independent estimates derived from a dual record system approach suggest that the discrepancies are generally quite moderate.

Results from our study confirm a trend toward lower infant and child mortality over the recent past as well as the existence of moderate regional mortality differentials and very substantial urban-rural mortality differences in Thailand. They also reveal substantial socioeconomic differentials in infant and child mortality and suggest that the overall level would be considerably reduced if the same health care were provided for the rural and socioeconomically disadvantaged populations as is apparently enjoyed by the numerically small, advantaged groups in urban areas.

A number of techniques to estimate infant and child mortality from census and survey data on children ever born and surviving children have been developed over the past decade (Brass et al., 1968; Sullivan, 1972; Trussell, 1975; Feeney, 1976). Details describing the actual calculations involved in applying several of these techniques have been conveniently summarized in recent handbooks (Lingner, 1974; Brass, 1975; National Academy of Sciences, Committee on Population and Demography, 1977). All of these techniques convert tabulations of the proportion dead among children ever born to women classified by standard age or marriage duration categories.

These indirect estimation techniques are particularly valuable in situations where reliable data on mortality are not available from conventional vital registration systems. Even when direct measures of mortality are available, indirect techniques are often useful for exploring differentials in mortality among subgroups in the population for which separate data are not provided in the conventional registration system. In the special case where mortality is being directly measured through a sample registration or dual record system, the indirect techniques provide a useful set of estimates for comparison. In addition, the recent developments in the indirect estimation approach now permit estimates of time trends in mortality that may not be so readily derived from sample registration or dual-record system efforts unless they have been in continuous operation over a fairly long period of time.

In Thailand, a vital registration system has been in operation since the early part of the century, but the data generated are generally considered incomplete. To estimate the extent of this incompleteness, as well as to arrive at accurate estimates of fertility and mortality, the government of Thailand conducted a Survey of Population Change (SPC) in 1964–67. A second SPC was started ten years later, in 1974, and a preliminary report of the results based on the first year of experience has already been published (National Statistical Office, 1977). Both Surveys of Population Change were based on a dual record system approach and yielded direct estimates of infant and child mortality. These results are presented in Table 1 along with the uncorrected infant mortality rate derived from the national registration system. Since there seems to be general agreement that for the earlier SPC the quality of the fieldwork during the first year of operation

TABLE 1 Direct measures of infant and child mortality in Thailand from uncorrected vital registration data and the Surveys of Population Change (SPC)

	Infant mortality rate[a]		Life table functions from SPC		
Period and sex	As registered	From SPC	$_1q_0$	$_4q_1$	$_5q_0$
1964–65					
Male	38.1	92.5	.0954	.0411	.1326
Female	30.6	76.0	.0753	.0398	.1121
Both sexes	34.5	84.3	.0856[b]	.0404[c]	.1226[b]
1974–75					
Male	30.5	70.7	.0919	.0309	.1199
Female	21.3	46.1	.0596	.0357	.0931
Both sexes	26.0	56.3	.0761[b]	.0333[c]	.1068[b]

a Deaths to infants under one year of age per thousand live births.

b Weighted by a sex ratio at birth of 1.05 males per 1.00 females.

c Weighted by a sex ratio at birth of 1.05 and adjusted for sex differences in $_1q_0$.

SOURCES: Ministry of Public Health (1972), Division of Vital Statistics (unpublished data), National Statistical Office (1969, 1977, 1978).

(which spanned 1964 and 1965) was superior to that during the subsequent two years, estimates based on the first year rather than the full three years are included in Table 1. Results for the later SPC also cover only the first year of operation since that is the period to which the preliminary report refers.

As can be seen from a comparison of the infant mortality rates derived from uncorrected registration data with those derived from the Surveys of Population Change, the registered rates suffer from severe degrees of incompleteness in both the 1964–65 period and the more recent period a decade later. Estimates of child mortality are available from the life tables calculated on the basis of the Surveys of Population Change. The results presented in Table 1 indicate that the probability of surviving between birth and age five increased slightly over the decade spanned by the two surveys. This increase in survival chances appears to be purely the result of an improvement in infant survival, for the probability of dying between age one and age five (${}_4q_1$) for the combined sexes actually shows a slight increase between the two surveys. Because the results are based on samples and there is always some uncertainty involved with estimates from the dual record system approach, comparisons between the two surveys should be treated with appropriate caution. In addition, the coverage of the two surveys differs somewhat. The second SPC covered the entire country whereas the first SPC omitted the Bangkok-Thonburi Metropolis. Furthermore, there are puzzling differences between the infant mortality rate and the life table function for infant mortality (${}_1q_0$) presented in the published results of the 1974–75 SPC. We suspect that the infant mortality rate is a more accurate reflection of actual mortality than the life table function since the central death rate under age one, on which the ${}_1q_0$ is based, may be inflated due to an underestimation of the base population under age one (Chamratrithirong, Debavalya, and Knodel, 1978).[1] Despite these problems of interpretation, the results from the Surveys of Population Change can serve as useful standards against which indirect estimates of infant and child mortality can be compared.

The purpose of the present study is to report estimates of infant

1 It is also possible, however, that the infant mortality rate underestimates the true mortality. It should be borne in mind that both births and deaths in the SPC are estimated using the Chandrasekaran-Deming formula (see National Statistical Office, 1977:68). If infant deaths were relatively underestimated compared with births, this would artificially depress the estimated infant mortality rate.

and child mortality derived through application of indirect estimation techniques to tabulations on children ever born and surviving children by age of women from the 2 percent sample of the 1970 census and the first-round results of the most recent Survey of Population Change. We will examine regional and urban-rural levels and trends as well as socioeconomic differentials in infant and child mortality. Before presenting our results in detail, however, we turn briefly to the choice of the specific techniques to be used in the subsequent analysis.

CHOICE OF TECHNIQUE

The original technique for estimating infant and child mortality from proportions dead among children ever born was developed by William Brass, and subsequent techniques have built upon it. The Brass technique and the Sullivan and Trussell modifications are all based on the assumption of constant fertility and mortality during the recent past, i.e., during the decade or two prior to the census or survey from which the data on proportions dead are drawn. Only the technique developed by Feeney is designed specifically to deal with changing mortality and yields time trends based on the assumption of a constant rate of change in mortality during the recent past. The other techniques, however, can be adapted to take declining mortality into account and also to yield time trends of mortality. Since the published description of the Feeney technique provides formulas for estimating only infant mortality, and not child mortality as do the other techniques, it is excluded from the following comparison.[2]

Table 2 compares the mortality estimates derived from the Brass, Sullivan, and Trussell techniques.[3] Since both the Sullivan and Trussell techniques are based on the Coale-Demeny model life tables (Coale and Demeny, 1966), the results depend on the choice of the particular family of life tables (North, West, South, and East) that is considered most appropriate for the population under investigation. Although the West model life tables are most commonly used, recent research by John Fulton (1975) has indicated that the North model life tables are

2 A computer program available at the East-West Population Institute can construct an entire life table based on the Feeney technique. The life table chosen is parameterized using the infant mortality rate. We did not have the means to compute the complete life table, however.

3 The Brass estimates are based on the use of P_1P_2 as the parameter for selecting the multiplier (see Brass, 1975:55). The Trussell estimates are based on a modified version of the coefficients published in Trussell (1975). The new coefficients are provided in Appendix Table 1.

TABLE 2 Indirect estimates of infant and child mortality calculated according to the Brass, Sullivan, and Trussell techniques as applied to the 1970 census and Round I of the 1974 Survey of Population Change: whole kingdom

Source and technique	$_1q_0$	$_2q_0$	$_3q_0$	$_5q_0$
1970 census				
Brass	.0713	.0790	.1019	.1249
Sullivan				
West model	a	.0783	.0990	.1202
North model	a	.0756	.0949	.1182
Trussell				
West model	.0689	.0782	.1006	.1242
North model	.0674	.0753	.0987	.1223
1974 SPC (Round I)				
Brass	.0734	.0709	.0831	.1077
Sullivan				
West model	a	.0703	.0808	.1038
North model	a	.0679	.0776	.1021
Trussell				
West model	.0666	.0695	.0822	.1077
North model	.0648	.0667	.0791	.1064

a Technique does not permit $_1q_0$ to be calculated.

SOURCE: Unpublished tables from the 2 percent sample of the 1970 census and Round I of the 1974 Survey of Population Change.

more appropriate for Thailand.[4] Results based on both the West and North models are shown in Table 2. Actually, the estimates are quite similar regardless of which technique or which model life table set is utilized. The largest difference is to be found between the Brass and Trussell estimates of the probability of surviving between birth and age one. (The Sullivan technique does not permit estimation of $_1q_0$.)

4 Fulton's analysis was based on the fit for the entire age schedule of mortality, whereas the fit at the early childhood ages is most relevant for the present study. A preliminary examination of the observed age pattern of child mortality as implied by a comparison of $_1q_0$ with $_4q_1$ taken from the 1964–65 and 1974–76 SPC results yields inconclusive results in part because of doubts concerning the accuracy of the SPC life tables. The data from the 1964–65 life table fit the West model better, while the data from the 1974–76 life table corresponded better to the North model. In the absence of more conclusive results we therefore decided to follow Fulton's findings and rely primarily on the North model for the main body of our analysis.

For both techniques, ${}_1q_0$ is estimated from the proportion dead among children ever born to women 15–19 years old, which renders the techniques particularly unreliable for several connected reasons. The total number of children born to women in these young ages is typically small and thus there is more random variability in the proportion dead derived from results for 15–19-year-old women than for older women. In addition, children born to women 15–19 are disproportionately first-order births which, in combination with the young age of the mothers, tend to exaggerate the level of infant mortality experienced by them, since both very young maternal age and first birth order are generally associated with higher than average mortality. Finally, estimates based on the 15–19 age group are very sensitive to the exact age pattern of early childbearing, which is relatively poorly accounted for by the methodology of the indirect estimation techniques.

Another way to compare the consistency of estimates derived by different methods is to compare the mortality levels, as expressed in the Coale-Demeny model life table system, that correspond to each set of estimates. In the Coale-Demeny model life table system, each successive level refers to an increase in female life expectancy at birth (e_0) of 2.5 years. Level 1 refers to a mortality level in which female e_0 is 20.0; level 2, to a female e_0 of 22.5; and so on until level 24, for which female e_0 is 77.5. Although the increase in male life expectancy between each successive level varies slightly, it is also approximately 2.5 years. Thus, for all practical purposes, an increase of one level for the combined sexes can be considered equivalent to an increase in e_0 of 2.5 years.

The model life table levels corresponding to the estimates of infant and child mortality shown in Table 2 are presented in Table 3.[5] The results show that the model life table levels for each of the infant or child mortality measures are quite similar for all three techniques regardless of whether the West or North model is used. Except for the ${}_1q_0$ estimates, the difference in model life table levels corresponding to a particular ${}_nq_0$ measure is rarely above four-tenths of a level. In other words, the mortality estimates derived from the three techniques generally correspond to mortality levels that differ by less than one

5 Although the results of the Brass technique are not based on the Coale-Demeny model life table and thus do not provide separate estimates for the North and West families, it is still possible to determine the corresponding Coale-Demeny life table level in each of the regional families that corresponds to each particular Brass estimate of infant and child mortality.

TABLE 3 Coale-Demeny life table levels corresponding to indirect estimates of infant and child mortality derived from the 1970 census and Round I of the 1974 Survey of Population Change, by regional model and estimation technique: whole kingdom

Source, regional model, and estimation technique	Model life table level corresponding to estimate of $_1q_0$	$_2q_0$	$_3q_0$	$_5q_0$
1970 census				
West model				
Brass	17.64	18.03	16.97	16.20
Sullivan	a	18.08	17.15	16.45
Trussell	17.86	18.08	17.05	16.24
North model				
Brass	17.10	17.72	16.82	16.38
Sullivan	a	17.98	17.27	16.75
Trussell	17.51	18.01	17.16	16.52
1974 SPC (Round I)				
West model				
Brass	17.46	18.61	18.14	17.13
Sullivan	a	18.66	18.29	17.35
Trussell	18.07	18.71	18.20	17.13
North model				
Brass	16.90	18.37	18.07	17.35
Sullivan	a	18.61	18.47	17.67
Trussell	17.78	18.71	18.36	17.42

a Technique does not permit $_1q_0$ to be calculated.

SOURCE: Same as for Table 2.

year in life expectancy at birth. Only the estimates of $_1q_0$ correspond to larger differences, although even in this case the difference is less than one full model life table level.

Assuming constant mortality in the recent past and assuming also that the data are accurate, we would expect the level of mortality corresponding to each of the different infant and child mortality estimates for any particular technique to be roughly similar. Instead, there is a clearly discernible pattern evident in Table 3. For all three techniques the mortality level declines as we move from the level corresponding to $_2q_0$ to the level corresponding to $_5q_0$. This pattern probably reflects recent declines in mortality. In each of these techniques the $_2q_0$ estimate is based on proportion of children dead among women aged 20–24, the $_3q_0$ estimate on proportion of children dead among women aged 25–29, and the $_5q_0$ estimate on the proportion of children dead among women of ages 30–34. Children born to women aged 30–34 were on the average born at an earlier period than children born to women aged 20–24. If mortality has been declining in the recent past, the children born to the older women will have been exposed to higher mortality levels than the children born to the younger women. The decline in the mortality levels between the estimates corresponding to the younger and older women indicates that indeed the children born to the older women were experiencing higher mortality (as indicated by lower model life table levels) and is consistent with a recent mortality decline. It seems unlikely that this pattern is due to reporting errors since it is reasonable to assume that the proportion dead would be less completely reported by the older women than the younger women because the older women are referring to experiences further in the past. If such a pattern of reporting errors does exist, it operates in a direction opposite to the pattern actually observed in Table 3 and moderates the differences in mortality levels that are observed.

Special mention should be made of the mortality levels corresponding to the $_1q_0$ estimates provided by the Brass and Trussell techniques. They do not fit into the overall pattern of declining levels with estimates based on women of increasing age. The model life table level associated with the $_1q_0$ estimates is typically lower (i.e., referring to a lower life expectancy) than are the levels corresponding to the $_2q_0$ estimates. This finding would seemingly contradict our assertion that declining mortality explains the differences in levels corresponding to the estimates of $_2q_0$ through $_5q_0$; but, as explained above, the $_1q_0$ estimates derived from these indirect techniques are typically inflated because they rely heavily on births to very young mothers and are disproportionately first-order births.

The main problem created by declining mortality for the indirect estimation of infant and child mortality based on the procedures described is that estimates derived from the proportion dead among children ever born to women of different ages refer to mortality prevailing at different times rather than the mortality of one particular period. Fortunately, it is now possible to handle this problem through a procedure that permits the estimation of the time period for the various mortality estimates that correspond to different age groups of women

(Coale and Trussell, 1977; National Academy of Sciences, Committee on Population and Demography, 1977). This procedure requires the assumption that mortality has been changing at a constant rate prior to the census or survey. It can be used for Thailand if we assume that the decline of mortality has been occurring in a linear fashion. By applying the appropriate regression equations for estimating the time for each separate estimate of mortality, we can find the average number of years in the past to which the average mortality level of our various estimates refers. (The regression equations for estimating this are reported in Appendix Table 2.) We will apply this technique in much of the subsequent analysis.

Since the mortality estimates derived from the Brass, Sullivan, and Trussell methods are very similar, we believe there is little point in presenting detailed results based on each of these methods in the subsequent analysis. Instead, we will rely on the Trussell technique, the most recently developed of the three, and will use the North model life table when applying the technique. Since the infant mortality estimate (as expressed by ${}_1q_0$) appears to be unreliable regardless of which technique is used, we will not estimate this directly from the proportion dead to women aged 15–19. Instead, we will determine the average mortality level within the Coale-Demeny system that corresponds to the estimates of ${}_2q_0$, ${}_3q_0$, and ${}_5q_0$ and then find the ${}_1q_0$ as well as the ${}_4q_1$ and e_0 from the appropriate North model life table.

An additional assumption of all the techniques under discussion is that fertility has been constant during the recent past. This assumption is necessary because the average parity of women in successive age groups at the time of the census or survey is taken to represent the experience of each cohort of women. There is considerable evidence that the assumption of constant fertility is not valid in the case of Thailand and that fertility has been declining since the mid or late 1960s. Although no adjustment will be made for this problem, it is possible to get a rough idea of how much it affects our estimates of child mortality from calculations described below. In brief, the effect seems to be minimal.

Coale and Trussell have developed a set of regression coefficients to be used in estimating child mortality from the proportion dead among children ever born when data on children ever born are available for *true* cohorts of women (Coale and Trussell, 1977; National Academy of Sciences, Committee on Population and Demography, 1977). The coefficients are provided in Appendix Table 3. If the actual cohort fertility experience of each five-year age group of women is known, we do not need to approximate it from the cross-sectional data provided in a single census or survey, and thus no assumption requiring constant fertility need be made. Such true cohort data for Thai women can be approximated by combining the results of the 1970 census and the 1974 Survey of Population Change.

If the interval between the collection of the first and the second of these data sets was exactly five years, one could represent cohort experiences by comparing women in each five-year age group in the census with women in the next older five-year age group in the SPC (ignoring for the moment that some sampling error is involved, since the SPC is based on a sample and the census data we are using are derived from tabulations from a 2 percent sample). In actuality, the interval between the census and the SPC is slightly less than four and one half years (the census enumeration took place on 1 April 1970 and the first round of the Survey of Population Change occurred in July through September 1974). Nevertheless, the interval is sufficiently close to five years that cohort fertility experience can be reasonably approximated. The following comparison, then, roughly indicates the impact of falling fertility on the estimates of Thai child mortality. (In both cases the proportion dead is taken from Round I of the 1974 SPC and the multipliers used to convert them into mortality measures are based on the North model coefficients.)

Approach used	${}_2q_0$	${}_3q_0$	${}_5q_0$
Using multipliers determined by "true cohort" data on children ever born (from comparing 1970 census and 1974 SPC)	.0654	.0767	.1027
Using Trussell multipliers determined by cross-sectional data on children ever born (from 1974 SPC only)	.0667	.0791	.1064
Ratio of the cohort-based estimate to the cross-sectional-based estimate	.98	.97	.97

These results suggest that the recent fertility changes in Thailand have only a small effect on the results obtained under the assumption of constant fertility.

As already noted, the data for the present study come from the 2 percent sample of the 1970 census and the first round of the second Survey of Population Change. In both cases the tabulations of children ever born and surviving children are from unpublished or prepublication tabulations made available by the National Statistical Office.[6] Since the data in the SPC were for ever married women only, we multiplied the average number of children ever born to ever married women by the proportion of women ever married to convert the average number of children into the required form for our calculations. Such a procedure ignores children born to single women, but in Thailand this is probably not a serious problem. We suspect that women who have borne a child are, by virtue of this fact, reported as ever married in most surveys.

NATIONAL, REGIONAL, AND URBAN-RURAL MORTALITY LEVELS

Results based on the Trussell technique (North model) are presented in Table 4 by urban-rural residence and regional location. The estimates of ${}_2q_0$, ${}_3q_0$, and ${}_5q_0$ shown are derived directly from the Trussell technique. Each of these estimates corresponds to a particular North model life table as well as to a particular time period that can be expressed as years prior to the census or survey (designated as t^*). It should be recalled that in circumstances of declining mortality, the ${}_2q_0$ estimate, which is based on women aged 20–24, will refer to a more

6 In the case of the Survey of Population Change, the published preliminary report includes tables indicating both children ever born and surviving children to ever married women on the basis of the first four rounds, which took place in 1974 and 1975. We have decided, however, to base our analysis on equivalent unpublished data from the first round only because we believe they are more appropriate for the application of the indirect techniques of mortality estimation described above. The published results on children ever born and surviving children covering all four rounds of the first year are based on all eligible respondents (i.e., ever married women) interviewed in the first round, plus all additional eligible respondents interviewed in each subsequent round. The final tabulation, then, is an average of the children ever born and surviving children reported by women interviewed in the first round and additional children reported by women interviewed for the first time in the next three rounds. We suspect that a disproportional number of the women interviewed in the second through fourth rounds were newly wed (hence explaining why they were not interviewed in the first round), although some were undoubtedly women who were already married at the time of the first round but could not be interviewed then due to a temporary absence or some related reason. If our assumption is correct, by including the women interviewed only after the first round, we would bias the results since the additional women will be disproportionally childless owing to the recency of their marriages. Thus we believe it is better to rely only on the first-round data, because they are more comparable to the usual cross-sectional data to which the indirect mortality estimation techniques are normally applied.

TABLE 4 Indirect estimates of child mortality, average mortality level, average time period of estimates (t^*), and corresponding $_1q_0$, $_4q_1$, and e_0 by urban-rural residence and region as derived from the 1970 census and Round I of the 1974 Survey of Population Change

Source and residence category	$_2q_0$	$_3q_0$	$_5q_0$	Average mortality level	Average t^* (years)	$_1q_0$, $_4q_1$, and e_0 corresponding to average level		
						$_1q_0$	$_4q_1$	e_0
1970 census (2 percent sample)								
Kingdom	.0753	.0967	.1223	17.23	4.0	.0701	.0426	58.7
Municipal areas	.0307	.0341	.0473	21.69	3.7	.0305	.0099	69.9
Nonmunicipal areas	.0799	.1029	.1309	16.82	4.0	.0742	.0462	57.7
Central region	.0604	.0726	.0922	18.78	3.8	.0554	.0300	62.6
North region	.0927	.1175	.1356	16.11	4.0	.0814	.0528	55.9
Northeast region	.0844	.1067	.1465	16.35	4.0	.0789	.0505	56.5
South region	.0484	.0810	.0963	18.85	4.3	.0547	.0295	62.7
1974 SPC (Round I)								
Kingdom	.0667	.0791	.1064	18.16	3.9	.0611	.0349	61.0
Municipal areas	.0280	.0280	.0408	22.17	3.4	.0269	.0133	71.1
Bangkok-Thonburi	.0177	.0139	.0264	23.59	3.5	.0170	.0034	74.8
Provincial urban	.0360	.0400	.0517	21.25	3.4	.0340	.0128	68.8
Nonmunicipal areas	.0704	.0843	.1135	17.80	4.0	.0646	.0378	60.1
Central region	.0304	.0402	.0673	21.08	3.8	.0354	.0139	68.4
Central region (excluding Bangkok-Thonburi)	.0327	.0458	.0763	20.65	3.9	.0390	.0167	67.3
North region	.0893	.0973	.1421	16.50	3.9	.0774	.0492	56.9
Northeast region	.0729	.0982	.1164	17.37	4.0	.0688	.0414	59.1
South region	.0810	.0673	.0955	18.28	4.2	.0600	.0339	61.3

NOTE: The estimates are derived by use of the Trussell technique, North model. The average mortality level and the corresponding $_1q_0$, $_4q_1$, and e_0 are all based on the Coale-Demeny North model life tables.

SOURCE: Same as for Table 2.

recent time period and presumably a higher model life table level (corresponding to lower mortality) than would be the case of the $_3q_0$ estimate, which is based on the children ever born to women five years older. Likewise, the $_5q_0$ estimate, which corresponds to the mortality experience of children to women aged 30–34, would refer to an even more distant time period and an even lower model life table level. The averages of the model life table levels implied by these three estimates of child mortality, as well as the average of the number of years prior to the census or survey to which the estimates refer, are included in Table 4. Finally, the values of $_1q_0$, $_4q_1$, and e_0 from the specific North model life table indicated by the average level are also presented. These last estimates thus refer to the average time period indicated, whereas the estimates derived directly from the Trussell technique correspond to different time periods as just described.

The age-gradient of mortality implied by the comparison of the three child mortality estimates obtained directly through the Trussell technique exaggerates the true age-gradient at any particular time if mortality has been declining prior to the date when the data were collected, because the $_2q_0$ estimate refers to the most recent time and the $_5q_0$ estimate refers to the most distant time. On the other hand, the values corresponding to the average level do not exaggerate the age-gradient of mortality since they simply reflect the age-gradient specified by the model life table corresponding to the average level.

Several features of Table 4 are worthy of note. First, the results from both the 1970 census and the 1974 SPC indicate a considerably lower level of child mortality in the urban sector (municipal areas) than in the rural sector (nonmunicipal areas). The Bangkok-Thonburi Metropolis, for which separate results are available from the 1974 SPC, is characterized by particularly low mortality. According to our results, the average life expectancy of a child born recently in Bangkok is almost 14 years more than for a child born in the rural areas. The overall rural-urban difference in life expectancy indicated by the 1974 SPC is approximately 11 years, compared with an urban-rural difference indicated by the 1970 census of about 12 years. Clearly, children born in the urban areas of Thailand today have considerably better survival chances than children born in the countryside. This urban advantage is undoubtedly due to a combination of higher socioeconomic levels in the cities and towns and the relatively easier access to superior medical and health services.

Second, there appear to be substantial regional differences in mortality. According to the results derived from the 1970 census, mortality conditions appear to be considerably worse in the north and northeast than in the south or central regions. The results from the 1974 SPC suggest that the advantages in the central region are not due simply to the inclusion of Bangkok since even if Bangkok is excluded the central region shows the most favorable regional mortality. Again in 1974, the north and northeast regions are characterized by the worst mortality conditions, although the northeast seems to be relatively better off than the north in more recent results. The south region occupies a position between the northeast and the central region according to the 1974 SPC and is actually characterized by slightly less favorable mortality than was indicated by the 1970 census results. It must be stressed that these results are only approximate, and the worsening of mortality in the south as indicated by the difference between the 1970 and 1974 results may reflect peculiarities of the data rather than an actual trend. Particularly in the case of the SPC, the results for the south are based on a relatively small number of cases. It is worth also recalling that the south is culturally more distinct from the rest of Thailand than any of the other regions and may be characterized by a different level of accuracy in the reporting of children ever born and surviving children, although we have no direct evidence to indicate this.

Third, the average time period to which the three mortality estimates refer is quite similar for each of the residence categories. In most cases, the average number of years prior to the survey for the three estimates is between four and four and one-half years. As noted above, the time period depends on the average age of the children ever born to the women in the age groups used for the mortality estimates. The children's average age in turn depends on the shape of the fertility schedule. Where the age at marriage is later, as for example in the municipal areas, the children born to women in their twenties are on the average younger than children born to women who married at earlier ages. This explains why the number of years prior to the census or survey to which the mortality estimates refer is shorter for municipal areas than for nonmunicipal areas. The unusually early age at marriage in the south relative to the rest of Thailand (Chamratrithirong, 1976) explains why the mortality estimates there refer to a particularly early period compared with the other regions.

Finally, there are apparent inconsistencies in the results derived from the first round of the 1974 SPC. For the municipal areas, the values of ${}_2q_0$ and ${}_3q_0$ are identical, whereas for the Bangkok-Thonburi Metropolis and the south region the values of ${}_2q_0$ actually exceed the values of $_3q_0$ according to the Trussell estimates. In reality, of course, the probability of dying before age two must be lower than the probability of dying before age three. Since our estimates of $_2q_0$, $_3q_0$, and $_5q_0$ are derived independently from each other, such inconsistencies can arise and probably result from a combination of small numbers of cases and reporting errors. Given the nature of the indirect estimation technique, however, it is rather remarkable that there are not more inconsistencies of this nature. In the three cases just pointed out, the number of respondents on which the estimates are based is particularly small. In addition, it is worth noting that a much higher proportion of the population of the south region is Muslim, while in all other regions the vast majority of the population is Buddhist. In a number of respects, the Muslim culture of the south is distinctive from the rest of Thailand and this difference could affect the extent of reporting errors, although there is no direct evidence of such an effect. In any event, all the figures in Table 4 should be treated as rough estimates rather than as precise indications of mortality levels.

One way to check the plausibility of the indirect estimates shown in Table 4 is to compare them with the direct measures of mortality derived from the Surveys of Population Change. Direct mortality measures by urban-rural residence and region from the first and second Surveys of Population Change are presented in Table 5. As indicated above, most official Thai reports utilize results based only on the first year of the 1964–67 SPC. Since no officially published life tables from either Survey of Population Change exist on a regional basis, we include in Table 5 unpublished results kindly provided by John Fulton from life tables by region based on the entire three years of the first Survey of Population Change.

Since the results for the indirect estimates of mortality based on the 1970 census refer on the average to a time period of about four and a half years prior to the census, it is appropriate to compare them with the direct measures obtained in the first SPC. The comparison is somewhat complicated since the results from the Survey of Population Change during the first year of operation do not agree completely with the results for the full three years of the survey. In general, however, the direct measures for either the first year or the first three years show higher mortality than do our indirect measures. For example, the infant mortality rate for 1964–65 as indicated by the SPC suggests that approximately 84 of 1,000 infants would die before their first birthday, the equivalent results for 1964–67 from the SPC indicate a 79 per thousand chance of dying before age one, whereas the $_1q_0$ estimate

TABLE 5 Direct measures of infant mortality and child mortality in the two Surveys of Population Change, by urban-rural residence and region

Region or residence	1964–65, IMR[a]	1964–67 $_1q_0$[b]	1964–67 $_4q_1$	1964–67 $_5q_0$	1964–67 e_0	1974–75, IMR[a]
Kingdom	84.3[c]	.079[c]	.050[c]	.125[c]	56.7[c]	56.3
Bangkok-Thonburi						31.0
Provincial urban	67.6					10.3
Nonmunicipal areas	85.5					63.9
Central region (excluding Bangkok-Thonburi)	94.0	.073	.030	.101	61.0	49.5
North region	96.5	.084	.050	.130	55.1	96.0
Northeast region	83.4	.085	.064	.144	53.2	54.4
South region	48.5	.059	.048	.104	61.0	60.4

a Infant mortality rate (IMR) is the number of infant deaths per thousand live births.

b For the purpose of calculating the 1964–67 life table, we assumed the population at age 0 to be equal to the number of births rather than deriving the age 0 population from the enumerated population under age 1. We did so because there was an obvious defect in the enumerated population data.

c Excluding Bangkok-Thonburi Metropolis.

SOURCE: National Statistical Office (1977); unpublished life tables calculated from the 1964–67 Survey of Population Change by John Fulton.

equivalent to the average of the indirect mortality estimates given in Table 4 suggests a 70 per thousand chance of dying before age one. This probably reflects an understatement of the proportion dead derived from the census questions on children ever born and surviving children. The indirect estimate of $_5q_0$ for the kingdom is also somewhat lower than the direct measures suggest (.122 versus .125), although the difference is not large and the indirect estimate includes the Bangkok-Thonburi Metropolis (which has lower than average mortality), while the direct estimate from the 1964–67 SPC does not. However, it should be borne in mind that the $_5q_0$ estimate is derived from the proportion dead among children ever born to women aged 30–34 and thus refers to an earlier time than the 1964–67 period covered by the SPC.[7]

7 The finding that there is better agreement between the direct and indirect $_5q_0$ estimates than for the $_1q_0$ estimates may reflect our choice of the North model life table as the appropriate model life table family for arriving at our estimates of $_1q_0$ and $_4q_1$. Of the four model life table sets, the North family is character-

In most cases, the indirect estimates of infant mortality for the regions are also lower than those indicated by the direct measures. One point of agreement between the indirect estimates and both sets of the direct measures is that infant mortality was lowest in the south region. Beyond this, however, the rank order of infant mortality in the remaining three regions differs for each of three sets of mortality figures. The indirect estimates suggest that the north region experienced the highest infant mortality; this inference agrees with the estimates for the 1964–65 Survey of Population Change results but differs from the 1964–67 results, which indicate that the northeast region has slightly higher infant mortality. In any event, all three sets of estimates indicate that the north region experienced above-average infant mortality. The biggest disagreement concerns the central region. It should be recalled that the first SPC excluded the Bangkok-Thonburi Metropolis from the central region whereas the 1970 census included it. The first-year results for the earlier Survey of Population Change suggest higher than average infant mortality in the central region (excluding Bangkok-Thonburi), whereas the 1964–67 results suggest slightly lower than average infant mortality in this area. The indirect estimates tend to agree more closely with the 1964–67 results, although they show an even more favorable picture than do the three-year results for the Survey of Population Change. This may be accounted for in part by the inclusion of Bangkok-Thonburi in the area covered by the indirect estimates. In addition, the 1974–75 SPC results indicate a considerably more favorable infant mortality rate in the central region than the other regions. In this case, then, the indirect estimates may be giving us a more accurate portrayal of regional differences during the mid 1960s than do the direct measures even though the indirect estimates probably underestimate the actual level of mortality for any particular region.

A comparison of the 1964–67 SPC estimates of ${}_5q_0$ with the indirect estimates derived from the 1970 census shows close agreement in both the rank order and the absolute level. Although the time period to which the indirect estimates refer is a few years earlier than the 1964–67 period, the agreement is still encouraging.

Regional differences in overall mortality, as indicated by the life expectancy corresponding to the indirect child mortality estimates from

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ized by the lowest values of ${}_1q_0$ relative to values of ${}_4q_1$. Thus we may be somewhat understating ${}_1q_0$ and overstating ${}_4q_1$ relative to each other if the true age gradient of mortality in the first five years of life in Thailand is more pronounced than implied by the North model life tables.

the 1970 census and the life expectancy calculated from the life tables for the 1964–67 Survey of Population Change, are in general agreement even if not completely consistent. In both cases, the south and central regions show similar and higher than average life expectancy whereas the north and northeast regions show similar but lower than average life expectancy.

A comparison between the ${}_1q_0$ implied by the indirect estimates of infant mortality based on the first round of the new Survey of Population Change with the direct measures of infant mortality based on the first four rounds suggests a fair amount of agreement. It should be remembered that the indirect estimates refer to a period four to five years earlier than that used for the direct estimates. In both cases, the lowest infant mortality is found for the central region and the highest mortality in the north region. In between, however, there is some disagreement between the estimates in the ordering of the northeast and south regions. The indirect estimate suggests that the south region has lower infant mortality than the northeast, whereas the direct estimates indicate the reverse. Considering that the estimates of infant mortality from the first SPC and the indirect estimates from the 1970 census substantially agree on the relatively low infant mortality in the south, it is not clear that the regional differentials as indicated by the direct measures are necessarily more valid than those obtained through the indirect measures.

Finally, it is worth noting that the direct measures of infant mortality from the 1974–75 Survey of Population Change suggest that infant mortality was higher in Bangkok than in the provincial urban areas. Moreover, the 10 per 1,000 mortality rate for the provincial urban areas is unreasonably low and would rival rates in the most advanced European country. In contrast, the indirect estimates from the proportion dead among children as reported in the first round of the Survey of Population Change indicate that infant mortality was lower in Bangkok than in the provincial urban areas, which seems inherently more plausible considering the superiority of health services in the capital. In addition, the higher level of infant mortality in provincial urban places seems more credible than the direct estimate.

In sum, comparison between direct and indirect estimates of regional and rural-urban differentials in infant and child mortality suggest some agreement. In general, the indirect estimates probably understate the mortality level to some extent. On the other hand, the relative mortality position of regions and of the rural and urban sections may be at least as accurately reflected in the indirect as in the direct measures.

TIME TRENDS IN INFANT MORTALITY

Feeney (1976) has proposed a technique for estimating time trends in infant mortality. Essentially, the technique derives estimates of $_1q_0$ from the proportion dead among children born to women of different ages as reported in a census or survey as well as an estimate of the number of years prior to the census or survey to which each of the infant mortality estimates refers.

As discussed in previous sections, Coale and Trussell have also developed a technique for estimating the number of years prior to the survey that indirect estimates of mortality based on the proportion dead among children ever born to women of different ages refer to. To estimate time trends in infant mortality from the indirect estimates of child mortality obtained through the Trussell technique, we have combined the Coale-Trussell estimates of years prior to the census or survey with the $_1q_0$ from the North model life table that corresponds to the level of mortality implied by the Trussell estimates of child mortality. For example, the Trussell technique yields an estimate of $_2q_0$ based on the proportion dead among children ever born to women aged 20–24. This $_2q_0$ corresponds to a particular level North model life table and thus to a particular $_1q_0$. This $_1q_0$ is then matched with the years prior to the census that correspond to the mortality experience of children born to women aged 20–24. A similar procedure is then followed in order to find the $_1q_0$ that corresponds to the $_3q_0$ derived from the proportion dead among children ever born to women aged 25–29, and to determine the number of years prior to the census or survey corresponding to the mortality experience of children born to women in this age group. Since the Trussell technique yields three estimates of child mortality (based on proportions dead among women of ages 20–24, 25–29, and 30–34) a time series of three estimates of infant mortality ($_1q_0$) can be derived in this manner. It should be noted, however, that the Feeney technique and the adaptation of the Trussell technique to yield time trend estimates of infant mortality are based on different underlying model life table systems and thus are not strictly comparable. The Trussell technique relies on the four-family model life table system developed by Coale and Demeny, whereas the Feeney technique is based on a one-parameter system developed by Brass.

The time trends in infant mortality based on both the Feeney technique and the adaptation of the Trussell technique are presented in Table 6 for the entire kingdom, the municipal and nonmunicipal areas separately, and the four regions as derived from the 1970 census, as

TABLE 6 Time trends in infant mortality based on indirect estimation survey to which each estimate corresponds, as applied to the

Technique used	Data from 1970 census (2 percent sample)					
	Kingdom		Municipal areas		Non-municipal areas	
	t^*	$_1q_0$	t^*	$_1q_0$	t^*	$_1q_0$
Feeney technique	2.4	.0584	2.1	.0228	2.4	.0622
	4.2	.0685	3.8	.0222	4.2	.0735
	6.4	.0768	5.8	.0277	6.4	.0828
	8.9	.0880	8.3	.0339	8.9	.0945
	11.9	.1011	11.2	.0417	11.9	.1080
	15.1	.1052	14.4	.0535	15.2	.1112
Adapted Trussell technique (North model)	2.2	.0625	2.2	.0281	2.2	.0659
	3.8	.0708	3.5	.0290	3.9	.0749
	6.0	.0772	5.3	.0346	6.1	.0820

well as for the entire kingdom as derived from the new Survey of Population Change. The trends for regions and for municipal and nonmunicipal areas based on the proportion dead from the new Survey of Population Change are not presented because the number of cases is quite small.

Both the Feeney technique and the adaptation of the Trussell technique confirm declining trends in infant mortality in all of the regional and rural-urban categories. Since the reasoning underlying both techniques is quite similar, it is not surprising that their results are also quite similar. It should be recalled that both techniques are based on an assumption of constant fertility in the recent past. This assumption is not strictly correct as applied to the Thai experience, particularly when we consider the estimates based on the 1974 data. As already discussed, the assumption of constant fertility is required because the children ever born to women of different ages found in cross-sectional data are assumed to represent the experience of a cohort of women passing through childbearing ages for the purpose of the calculation involved in the indirect mortality estimates as well as the calculation of the time period of the estimates. Since the 1970 census and the first round of the new Survey of Population Change were almost five years apart, by combining the data on children ever born from these techniques of ${}_1q_0$ and the number of years (t^*) prior to the census or 1970 census and Round I of the 1974 Survey of Population Change

Central region		North region		Northeast region		South region		Data from 1974 SPC (Round I): kingdom	
t^*	${}_1q_0$	t^*	${}_1q_0$	t^*	${}_1q_0$	t^*	${}_1q_0$	t^*	${}_1q_0$
2.1	.0471	2.4	.0721	2.1	.0670	2.5	.0383	2.2	.0532
3.9	.0501	4.2	.0848	3.9	.0783	4.3	.0571	3.9	.0561
5.7	.0560	6.4	.0862	6.0	.0957	6.5	.0591	6.0	.0665
8.5	.0596	8.9	.1010	8.5	.1101	9.0	.0659	8.5	.0707
11.4	.0747	11.9	.1220	11.5	.1187	12.1	.0795	11.4	.0807
14.6	.0824	15.2	.1304	14.7	.1185	15.3	.0785	14.7	.0890
2.2	.0512	2.1	.0755	2.0	.0694	2.6	.0419	2.3	.0560
3.5	.0549	3.9	.0842	4.0	.0773	4.0	.0605	3.8	.0593
5.8	.0602	6.0	.0816	6.0	.0906	6.2	.0625	5.7	.0683

two sources we can approximate the experiences of actual cohorts of five-year age groups of women.

We have already compared the results of the mortality estimates obtained by using the cohort experience with the estimates obtained by using the cross-sectional data and found that the difference is rather minimal. We can make the same comparison to see how much the violation of the assumption of constant fertility affects our time-trend estimates derived from the adapted Trussell technique. In other words, we can substitute data on children ever born to actual cohorts of women in the formula for the calculation of the years prior to the census to which the estimates refer. We can also find in the North model life table the ${}_1q_0$ value that corresponds to the indirect mortality estimates obtained by using multipliers determined from "true cohort" data on children ever born. We can compare these results with the equivalent data obtained using cross-sectional data on children ever born. The following comparison, then, roughly indicates the effect of falling fertility on the time-trend estimates of infant mortality obtained by the adapted Trussell technique.[8]

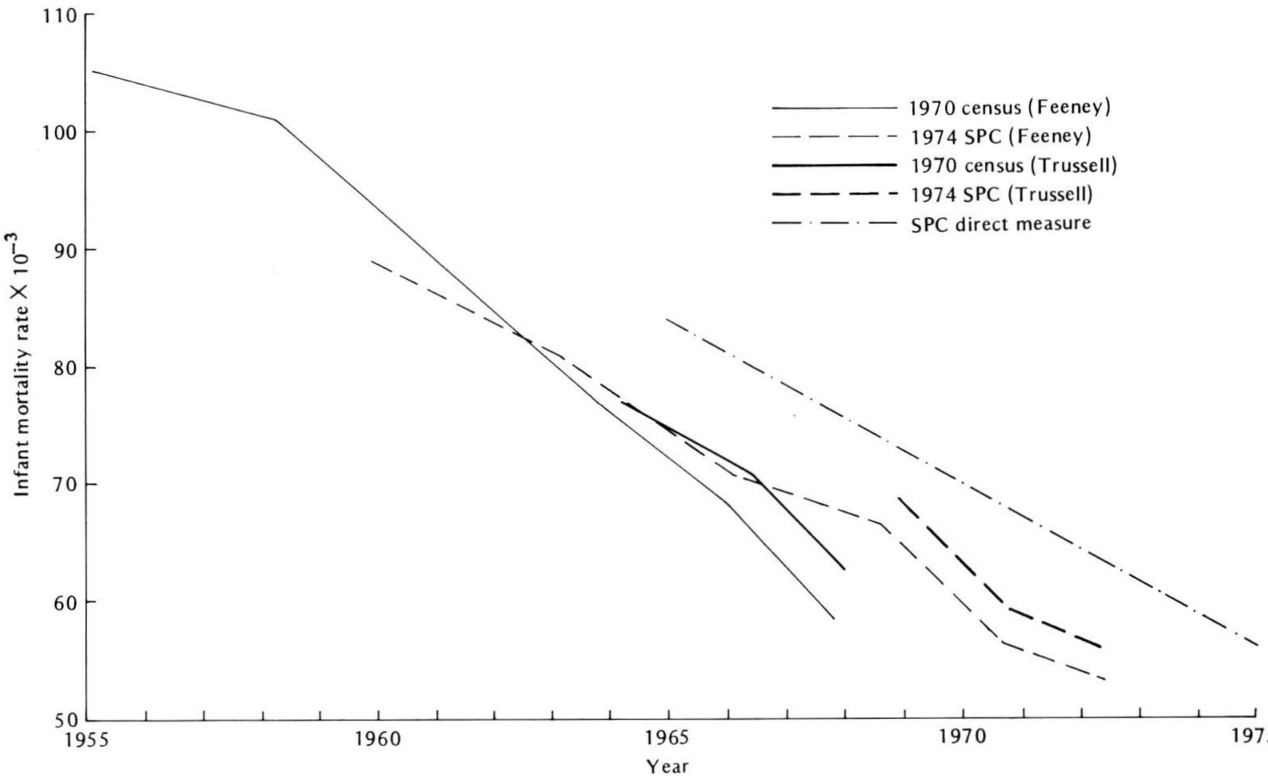
8 In both cases the proportion dead, $D(j)$, is taken from Round I of the 1974 SPC; for the estimates of t^* using cohort data $P(j)$ is taken from the 1970 census and $P(m)$ from Round I of the 1974 SPC; for the estimate of t^* using cross-sectional data both $P(j)$ and $P(m)$ are taken from Round I of the 1974 SPC.

Using cohort data on children ever born		Using cross-sectional data on children ever born	
t^*	${}_1q_0$	t^*	${}_1q_0$
2.3	.0550	2.3	.0560
4.1	.0570	3.8	.0593
6.5	.0662	5.7	.0683

As can be seen from the text table, the time trends are reasonably similar whether we use actual cohort data or use cross-sectional data for the children ever born. The actual cohort data result in placing the time period of the estimates slightly further in the past, at least for the estimates based on the proportion dead among children born to women aged 25–29 and women 30–34. Moreover, the ${}_1q_0$ values are slightly lower when cohort data are used. Since infant mortality appears to be declining, both of these differences reinforce each other in a direction that would lead estimates based on cross-sectional data to overestimate infant mortality at any particular time prior to the census or survey. We have already observed, however, that the indirect mortality estimates, when based on cross-sectional data, actually tend to underestimate mortality, as seen from comparisons with direct measures of mortality which presumably are more accurate reflections of the true level of mortality. Thus the bias created by using cross-sectional fertility data when fertility is actually declining appears in the case of Thailand to compensate in part for an even larger error in the opposite direction, which results from an apparent understatement of the proportion of children dead as indicated by comparisons of reported surviving children with reported children ever born.

The trend estimates of infant mortality presented in Table 6 are shown graphically in Figures 1, 2, and 3 along with the equivalent direct estimates of infant mortality based on the two Surveys of Population Change. Figure 1 plots the various estimates of trends in infant mortality for the whole kingdom. Although the levels indicated by the indirect estimates (whether based on the Feeney technique or the adapted Trussell technique) are below the levels indicated by the direct measures derived from the Surveys of Population Change, the general trends indicated by both the direct and the indirect measures are reasonably parallel. In addition, there is partial but not perfect correspondence between the time series of indirect estimates based on the 1970 census and the 1974 Survey of Population Change. If we focus on the Feeney estimates, we can see that there is reasonable overlap in the

FIGURE 1 Trends in infant mortality for the whole kingdom



two series up to about 1966 but then there is some divergence. The same divergence is also implicit in the Trussell estimates, although the time series of the Trussell estimates do not actually overlap in the time period to which they refer. Perhaps the most important point to note, however, is that both the series of indirect estimates of infant mortality and the direct estimates lead to the conclusion that infant mortality has been declining in Thailand, at least over the last decade or two.

The rate of decline in infant mortality can also be estimated from the time series shown in Table 6. Feeney (1976) has suggested that the rate of mortality decline indicated by the estimates from a single census may reflect such spurious elements as relatively high mortality for the children of young mothers or under-reporting of deceased children. He therefore recommends estimating the rate of decline by comparing time series from two successive censuses or surveys. We have derived estimates of the rate of infant mortality decline by comparing the infant mortality estimates for each pair of corresponding age groups in the 1970 census and the 1974 Survey of Population Change. Feeney suggests that such a comparison is superior since errors and biases may be similar for particular age groups in each census or survey. For example, we arrive at the rate of decline in infant mortality implied by the proportion dead among children born to women aged 20–24 by subtracting the infant mortality rate associated with this age group in the 1974 Survey of Population Change from the infant mortality rate associated with the same age group in the 1970 census and then divide this difference by the number of years that separate the time periods associated with each of these estimates. By doing this for each age group, we arrive at the results presented in Table 7.

The irregularity of the estimates can be dealt with by aggregating them over all age groups. Thus the average annual rate of decline is shown for both the Feeney technique and the adapted Trussell technique. In the former case the average rate of decline is approximately 2.8 infants deaths per thousand live births per year, while in the latter case it is approximately 2 infant deaths per thousand per year. It is worth noting that the rate of decline implied by the direct estimates of infant mortality from the 1964–65 and the 1974–75 Surveys of Population Change is 2.8 infant deaths per thousand live births.[9] Thus,

9 This figure is based on a comparison of infant mortality rates (which relate infant deaths to live births) given for the Survey of Population Change rather than the ${}_1q_0$ estimates from the life tables. If we use the ${}_1q_0$ estimates, the rate of decline would only be 1.0 per thousand. As discussed previously, we have reason to believe that the infant mortality rates for the 1974–75 Survey of

TABLE 7 Annual rate of decline in infant deaths per 1,000 live births implied by Feeney technique and adaptation of Trussell technique

Age group of mothers	Feeney technique	Adapted Trussell technique
20–24	1.1	1.5
25–29	2.6	2.6
30–34	2.1	1.9
35–39	3.6	na
40–44	4.2	na
45–49	3.4	na
Average	2.8	2.0

na—not applicable.

despite the higher level of infant mortality indicated by the Surveys of Population Change, the rate of decline is fairly similar to that implied by the indirect estimates.

Figure 2 shows graphically the trends in the indirect estimates of infant mortality in municipal and nonmunicipal areas based on the 1970 census as well as the direct estimates based on the SPC for nonmunicipal areas. (The direct SPC estimates for municipal areas are not shown because the first Survey of Population Change excluded Bangkok-Thonburi from the sample and thus there are no equivalent estimates of infant mortality in municipal areas from that survey.) The dramatic difference in infant mortality rates in the municipal and nonmunicipal areas is clearly evident in Figure 2. Nevertheless, for both rural and urban sectors of the population, infant mortality appears to have been on the decline in the recent past. Indeed, the trend seems to be declining at least as steeply in the nonmunicipal areas as in the municipal areas.

Figure 3 plots the indirect estimates of the time series of infant mortality by region. The contrast between the higher infant mortality in the north and northeast compared with the central region and the south is quite evident. On the other hand, the decline in infant mortality appears to be taking place in all four regions and is at least as steep in the north and northeast as in the central and south regions.

Population Change is a better index of mortality than the ${}_1q_0$ figure from the life table since the latter is based on the central death rate under age one, which may need some correction.

FIGURE 2 Trends in infant mortality for municipal and nonmunicipal areas

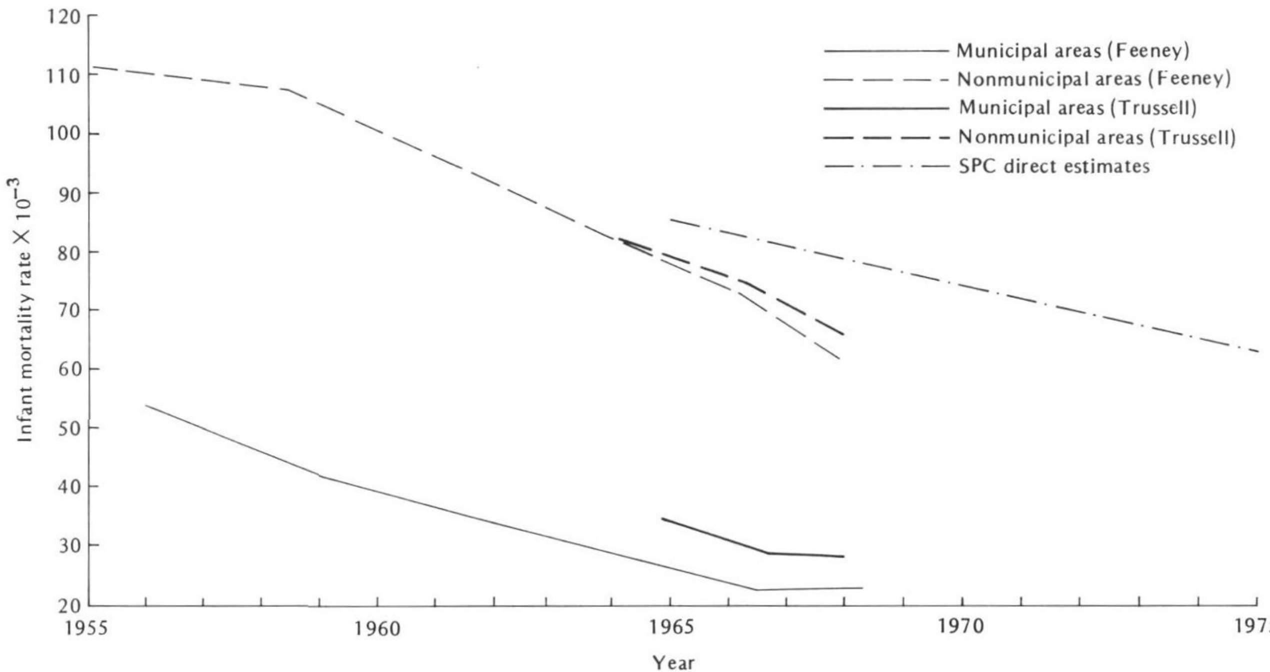
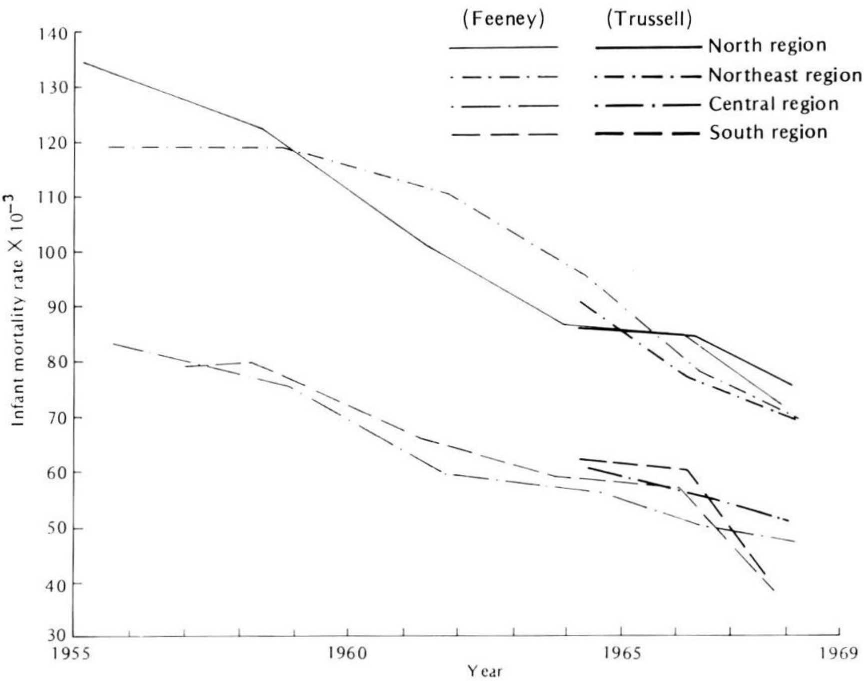


FIGURE 3 Trends in infant mortality, by region



In sum, the level of infant mortality indicated by indirect estimation techniques is probably understated, but the trends indicated by these techniques appear to be reasonable and are more or less consistent with the direct estimates of changes in infant mortality over the recent past. In Thailand, infant mortality appears to have been declining not only at the national level but also at the regional and rural and urban levels.

SOCIOECONOMIC DIFFERENTIALS IN MORTALITY LEVELS

We have already examined differentials in infant and child mortality based on place of residence as defined by regional location for urban-rural residents. Special tabulations from the 2 percent sample of the 1970 census permit us to estimate infant and child mortality through the indirect techniques according to selected socioeconomic characteristics of the mother. In this section we focus on differentials in mortality according to educational attainment, religion, labor force participation, and occupation of the mother.[10] To control for the substantial urban-rural differences in infant and child mortality, we will present results separately for municipal and nonmunicipal areas as well as for the whole kingdom standardized for urban-rural residence.

The important influence of the mother's educational attainment on the infant's or child's chances for survival are shown in Table 8. In both municipal and nonmunicipal areas as well as in the kingdom as a whole, there is a sharp inverse relationship between the level of schooling achieved by the mother and the mortality experience of the child. Indeed, the probability of dying before age two, three, or five is at least several times higher in both nonmunicipal and municipal areas for children of mothers who have had no schooling than for children of mothers who have had at least a secondary level education. For each measure, mortality decreases consistently with the educational attainment of the mother, although the difference is more pronounced between an elementary education (1–7 years) and secondary or university education than between no schooling and an elementary education. The most extreme differences are between children born to rural mothers with no schooling and children born to urban mothers with at least a secondary education. The chances of dying before age five are seven times greater for the former than for the latter.

The unstandardized differentials for the kingdom as a whole are even more pronounced than they are for either the nonmunicipal or municipal areas because of the association between higher education and urban residence. Women with no schooling are disproportionately rural while women with a secondary school education or above are disproportionately urban. In the first case both factors are associated

10 Unlike educational attainment or religion, which are basically permanent attributes of women by the time they reach the reproductive years, occupation and labor force participation may change over time. Thus the assumption implicit in the indirect mortality estimation approach that the average parity for successive age groups of women in a particular occupational or labor force category reflects the past fertility experience of each cohort is not likely to be correct. (This problem is independent of that of changing fertility during the recent past.) The age gradient of fertility for the different occupational and labor force participation categories of women may therefore be inaccurately estimated by the values of $P_{(1)}$, $P_{(2)}$, and $P_{(3)}$. A similar problem may characterize our estimates of rural-urban differentials since residence can also change during the childbearing years. The extent to which this problem will affect our estimates is unknown but it is unlikely to be very large.

TABLE 8 Education: indirect estimates of child mortality, average mortality level, average time period of estimates (t^*), and corresponding ${}_1q_0$, ${}_4q_1$, and e_0 as derived from the 1970 census, by educational attainment of mother, urban-rural residence, and (for whole kingdom) unstandardized and standardized rates for urban-rural residence

Residence and educational attainment	${}_2q_0$	${}_3q_0$	${}_5q_0$	Average mortality level	Average t^* (years)	${}_1q_0$, ${}_4q_1$, and e_0 corresponding to average level: ${}_1q_0$	${}_4q_1$	e_0
Nonmunicipal areas								
No schooling	.0975	.1249	.1515	15.57	4.5	.0870	.0581	54.6
Elementary	.0768	.0986	.1250	17.10	4.0	.0714	.0437	58.2
Secondary or university	.0284	.0278	.0416	22.13	3.3	.0271	.0075	71.0
Municipal areas								
No schooling	.0474	.0421	.0545	20.76	4.1	.0380	.0160	67.6
Elementary	.0267	.0361	.0492	21.75	4.0	.0301	.0095	70.1
Secondary or university	.0176	.0162	.0205	23.72	3.3	.0161	.0031	75.2
Kingdom								
Unstandardized								
No schooling	.0946	.1183	.1418	15.94	4.5	.0831	.0544	55.5
Elementary	.0733	.0931	.1181	17.44	4.0	.0681	.0408	59.3
Secondary or university	.0228	.0208	.0285	23.01	3.2	.0209	.0047	73.3
Standardized								
No schooling	.0927	.1170	.1422			.0823	.0541	55.8
Elementary	.0720	.0926	.1176			.0674	.0404	59.3
Secondary or university	.0274	.0267	.0396			.0260	.0071	71.4

NOTE: The estimates are derived by use of the Trussell technique, North model. The average mortality level and the corresponding ${}_1q_0$, ${}_4q_1$, and e_0 are all based on the Coale-Demeny North model life tables. The standardized figures result from assigning weights of .096 to the municipal area results and .904 to the nonmunicipal area results. These weights correspond to the urban and rural distribution of births to women in all educational categories in the five years prior to the 1970 census as estimated by reverse survival of the population aged 0–4 in municipal and nonmunicipal areas. The reverse survival procedure was based on the Coale-Demeny North model life table corresponding to the levels indicated in Table 4 from results based on the 1970 census.

SOURCE: Unpublished tables from the 2 percent sample of the 1970 census.

with higher mortality, whereas in the second case both factors are associated with lower mortality and thus urban-rural residence and education tend to reinforce each other in their impact on child mortality at the national level. National-level differentials in child mortality by the educational attainment of the mother are also presented in standardized form for urban-rural residence. The standardization was achieved by weighting the separate municipal and nonmunicipal area results for each educational category by the overall rural-urban distribution of births in the five years prior to the 1970 census for women of all educational categories. Although the standardization had the effect of contracting the differentials compared with the unstandardized results, substantial differences still exist in the mortality of children born to women of different educational levels. For example, the probability of dying before age five is three and a half times greater for children born to mothers who have had no schooling than for children born to mothers with a secondary or university education, even after standardization for urban-rural residence.

Not only is the educational attainment of the mother associated with the level of child mortality after urban-rural residence has been controlled, but also urban-rural residence continues to be strongly associated with mortality after the educational attainment of the mother has been controlled. Mortality risks are consistently higher for children born in nonmunicipal areas than for those born in municipal areas for each of the three categories of educational attainment of mothers. In relative terms, the urban-rural differences are substantial for all three educational categories, although in absolute terms the differences are much less for the higher educational group because in both nonmunicipal and municipal areas the mortality of children to better educated women is rather low.

Although the results in Table 8 may underestimate the level of infant and child mortality because they are based on the indirect estimation procedures, there is no apparent reason to suspect that the differentials indicated are not genuine. In fact, we might reasonably expect that women with less education would be more likely than women with higher education to omit deceased children from their reports of children ever born. If this were so, it would disproportionately bias the estimates of mortality of women in the lower educational categories downward compared with women in the higher educational categories and thus operate in a direction opposite to the mortality differentials actually observed. Several inconsistencies in Table 8 should be noted. For the women in the highest educational category in both municipal and nonmunicipal areas and for women with no schooling in municipal areas, the estimate of ${}_3q_0$ is lower than the estimate of ${}_2q_0$. Obviously these estimates do not accurately reflect the actual situation. Although without additional information it is impossible to say definitively what accounts for the discrepancy, the small number of women on which the results are based for these categories undoubtedly detracts from the reliability of the results. In addition, women with secondary or university education marry later than those with lesser education (Institute of Population Studies and National Statistical Office, 1977:44). Since many of these women marry in their early twenties, the ${}_2q_0$ estimate, which is based on the proportion dead among children ever born to women aged 20–24, may be heavily weighted by the mortality experience of first-born children, thus inflating the ${}_2q_0$ estimate. This discrepancy, of course, does not contribute to the inconsistency observed for urban women with no schooling.

Mortality estimates by religious affiliation are presented in Table 9. The Thai population is overwhelmingly Buddhist, with 95 percent of Thais classified as such in the 1970 census. Moslems constitute only 4 percent of the population whereas all other religious groups together represent less than 1 percent. Because of the very small number of cases available for groups other than Buddhists or Moslems, results are presented for only those two groups. Even for Moslems, the number of cases involved is small and the results for them need to be treated with additional caution.

For the kingdom as a whole as well as for the rural population there appears to be little difference between Buddhist and Moslem children in their chances of dying. Among the urban population, however, Moslem children appear to experience distinctly higher mortality than Buddhist children. Without additional information on the characteristics of each of the religious groups it is difficult to interpret this finding. We do know from the census that about the same proportion of urban Buddhists and urban Moslems live in the Bangkok-Thonburi Metropolis, so this in itself cannot play a part in accounting for the differential. Since the number of cases on which the mortality estimates for urban Moslems are based is extremely small, it is possible that the differentials observed may simply reflect sampling error. At this point we must leave the question unresolved.

Standardizing the results for urban-rural residence as we did in the previous table has little impact on the results because the distribution of the two religious groups with respect to urban-rural residence is quite similar. For both Buddhists and Moslems, however, consistently

TABLE 9 Religion: indirect estimates of child mortality, average mortality level, average time period of estimates (t^*), and corresponding ${}_1q_0$, ${}_4q_1$, and e_0 as derived from the 1970 census, by religion of mother, urban-rural residence, and (for whole kingdom) unstandardized and standardized rates for urban-rural residence

Residence and religion	${}_2q_0$	${}_3q_0$	${}_5q_0$	Average mortality level	Average t^* (years)	${}_1q_0$, ${}_4q_0$, and e_0 corresponding to average level		
						${}_1q_0$	${}_4q_1$	e_0
Nonmunicipal areas								
Buddhism	.0802	.1023	.1310	16.82	4.0	.0742	.0462	57.7
Islam	.0728	.1070	.1348	16.85	4.6	.0739	.0459	57.8
Municipal areas								
Buddhism	.0289	.0323	.0455	21.86	3.6	.0292	.0088	70.3
Islam	.0646	.0748	.0903	18.65	4.2	.0566	.0310	62.2
Kingdom								
Unstandardized								
Buddhism	.0749	.0953	.1222	17.27	4.0	.0697	.0423	58.8
Islam	.0715	.1046	.1312	17.00	4.6	.0724	.0445	58.1
Standardized								
Buddhism	.0753	.0956	.1228			.0699	.0426	58.9
Islam	.0720	.1039	.1305			.0722	.0445	58.2

NOTE: See note for Table 8.

SOURCE: Same as for Table 8.

higher mortality is evident for the rural population for all the mortality measures, although the contrast is less pronounced for the Moslems owing to their relatively high urban mortality.

In Thailand, female labor force participation is relatively high. For example, according to the 1970 census, among women in the ages 20–34 (whose children form the basis of the indirect mortality estimates employed in the present study), 86.4 percent in the rural areas, 51 percent in the urban areas, and 81 percent in the kingdom as a whole were in the labor force. The higher labor force participation among rural women is attributable to the inclusion of unpaid women working in family farming enterprises in the definition of the labor force incorporated in the Thai census. Indeed, 91 percent of rural women aged 20–34 in the labor force were in the occupational category of "farmers and miners." Because the Thai population is predominantly rural and rural women in the labor force are overwhelmingly in one occupational category, the number of women in the 2 percent census sample in other occupational categories is typically small and detracts from the reliability of our estimates of infant and child mortality by mother's occupation, which are presented in addition to differentials by labor force status in Table 9. Some of the unreliability can be overcome by focusing on the average mortality level indicated by the several measures yielded by the Trussell technique rather than taking each individual measure literally.

The results presented in Table 10 indicate that children of mothers who are in the labor force experience higher risks of dying in infancy and childhood than do children of mothers not in the labor force. This finding holds true for all the mortality measures shown in both rural and urban areas. It is difficult, however, to interpret the finding, for several reasons. Although it is possible that labor force activity competes with child care and breast feeding, which would have the effect of increasing mortality and thereby directly contribute to the observed differential, indirect relationships also seem plausible. For example, women who remain outside the labor force might be financially better off on the average than women who find it necessary to work and the lower mortality of their children might simply reflect this difference. In addition, interpretation is further complicated by the possibility that the mortality experience of a woman's children could affect her likelihood of working. Unlike educational attainment and religion, which are characteristics that are largely determined prior to childbearing and thus are unlikely to have been influenced by the mortality experience of a woman's children, labor force status could be so influ-

TABLE 10 Occupation: indirect estimates of child mortality, average mortality level, average time period of estimates (t^*), and corresponding $_1q_0$, $_4q_1$, and e_0 as derived from the 1970 census, by labor force participation of mother, occupation, urban-rural residence, and (for whole kingdom) unstandardized and standardized rates for urban-rural residence

Residence, labor force status, and occupation	$_2q_0$	$_3q_0$	$_5q_0$	Average mortality level	Average t^* (years)	$_1q_0$, $_4q_1$, and e_0 corresponding to average level		
						$_1q_0$	$_4q_1$	e_0
Nonmunicipal areas								
Outside labor force	.0655	.0909	.1204	17.66	4.3	.0661	.0390	59.8
In labor force	.0829	.1053	.1354	16.66	4.0	.0758	.0477	57.3
Professional, administrative, clerical	.0804	.0211	.0635	20.34	3.4	.0416	.0188	66.5
Sales	.1212	.0754	.1202	16.65	3.8	.0759	.0478	57.3
Farming	.0825	.1067	.1333	16.63	4.0	.0761	.0480	57.2
Other	.0660	.0975	.1309	17.31	3.5	.0693	.0419	58.9
Municipal areas								
Outside labor force	.0291	.0329	.0450	21.84	3.9	.0293	.0089	70.3
In labor force	.0344	.0362	.0514	21.42	3.3	.0327	.0117	69.2
Professional, administrative, clerical	.0058	.0155	.0278	24.00	3.7	.0142	.0024	75.9
Sales	.0242	.0373	.0500	21.79	3.4	.0297	.0093	70.1
Farming	.0634	.0633	.0988	18.79	3.9	.0553	.0300	62.6
Other	.0443	.0385	.0496	21.07	3.2	.0355	.0139	68.3
Kingdom								
Unstandardized								
Outside labor force	.0548	.0743	.0986	18.77	4.1	.0555	.0301	62.5
In labor force	.0810	.1022	.1279	17.06	3.9	.0718	.0440	58.3
Professional, administrative, clerical	.0368	.0175	.0421	22.24	3.6	.0264	.0071	71.3
Sales	.0853	.0624	.0923	18.36	3.6	.0593	.0333	61.5
Farming	.0824	.1065	.1330	16.64	4.0	.0760	.0479	57.2
Other	.0582	.0729	.0955	18.76	3.4	.0555	.0302	62.6

Standardized						
Outside labor force	.0620	.0853	.1132	.0626	.0361	60.8
In labor force	.0782	.0987	.1273	.0717	.0442	58.5
Professional, administrative, clerical	.0732	.0206	.0601	.0390	.0172	67.4
Sales	.1119	.0717	.1135	.0715	.0441	58.5
Farming	.0807	.1025	.1300	.0741	.0463	57.8
Other	.0639	.0918	.1231	.0661	.0392	59.8

NOTE: The farming category includes a small number of miners; the "other" category includes labor, crafts, transport, service, and production. See also note for Table 8.

SOURCES: Same as for Table 8.

enced. For example, a woman whose children have died may have freer use of her time and thus be more likely to work than a woman whose children survived. Our data, of course, do not permit us to determine if any such relationship operates in this direction, but the possibility should be kept in mind.

The influence of urban-rural residence remains strong in both categories of labor force participation. All the mortality measures consistently indicate that the children of working urban mothers had higher chances of survival than children of nonworking rural mothers. On the kingdom level, standardization for urban-rural residence reduces the differential in infant and child mortality by the labor force status of the mother, because the association of higher mortality with rural residence and labor force participation is mutually reinforcing for the unstandardized rates.

Mortality differentials according to the mother's occupation are somewhat unstable, probably because of the small number of cases in several of the occupational categories. In both rural and urban areas, however, the most favorable mortality levels clearly characterize the children of women who are in professional, administrative, and clerical occupations, a finding that undoubtedly reflects the advantages of their favorable socioeconomic position. The worst mortality experience characterizes children of farming women, although in rural areas the results are not consistent across the three separately derived mortality measures. In general, urban-rural differences in mortality seem to hold across all occupational categories.

In sum, our examination of differentials in infant and child mortality in Thailand based on indirect measurement reveals a sharp contrast between the socioeconomically most favored and the least favored groups. The most pervasive difference is an urban-rural one that seems to hold across every social and economic grouping. Within both the rural and urban populations, however, social class differences as measured by educational level or occupation of mother are also important. Although the level of mortality indicated by our measures may not be precisely accurate, we have no reason to believe that the differentials indicated are spurious. If we can further accept the levels indicated as not too far off the mark, the results suggest that the children of the more privileged strata in Thailand enjoy survival chances as favorable as those characterizing populations in the most economically advanced nations. Our results suggest that even a decade ago children born to urban mothers with secondary or university education, or to urban mothers employed in professional, administrative, and clerical occupations, experienced infant or child mortality levels that imply life expectancies of 75 years or more. Even in the rural areas, the mortality of children of the best educated mothers implies a life expectancy of 71 years. In contrast, the majority of children in Thailand apparently fare considerably worse.

Although all the indirect estimates may tend to be too optimistic because of under-reporting of dead children in the census and surveys on which the present study is based, comparison with direct measures suggests they are not too far off the mark. Furthermore, if mortality has improved since the time period covered by these estimates, life expectancies in excess of 75 years may not be unrealistic at present for the higher socioeconomic strata. Thus, within Thailand, the conditions to reduce mortality to extremely favorable levels are already present for certain segments of society. The real challenge is to create those conditions for a far broader spectrum of the population.

CONCLUSIONS

The application of indirect estimation techniques to determine infant and child mortality from survey and census data on the proportion dead among children ever born in Thailand demonstrates their usefulness in supplementing mortality data derived from direct measurement approaches. Comparison of results based on the direct and indirect techniques suggests that although the latter may underestimate mortality to some extent, the discrepancies are generally quite moderate. The indirect measurement techniques have the advantage of providing estimates of socioeconomic differentials and of time trends in infant mortality that would be difficult and extremely expensive to obtain through such procedures as sample registration or the dual record approach.

Results for our study confirm a trend toward lower infant and child mortality over the recent past as well as the existence of moderate regional mortality differences and very substantial urban-rural mortality differences in Thailand. In addition, they provide estimates for the first time of socioeconomic differentials in mortality and reveal large gaps between the survival chances experienced by children in the numerically small higher socioeconomic groups and the survival chances for children of the remainder of the population. Documentation of such differences should help to guide a national health policy.

APPENDIX TABLE 1 Regression coefficients (for North and West mortality patterns) to be used in estimating multipliers for the Trussell version where children ever born are classified by age of mother

Coale-Demeny mortality pattern and age group	$q(x)/D(j)$	$A(i)$	$B(i)$	$C(i)$
North				
15–19	$q(1)/D(1)$	1.1119	−2.9287	.8507
20–24	$q(2)/D(2)$	1.2390	−0.6865	−.2745
25–29	$q(3)/D(3)$	1.1884	0.0421	−.5156
30–34	$q(5)/D(4)$	1.2046	0.3037	−.5656
West				
15–19	$q(1)/D(1)$	1.1415	−2.7070	.7663
20–24	$q(2)/D(2)$	1.2563	−.5381	−.2637
25–29	$q(3)/D(3)$	1.1851	.0633	−.4177
30–34	$q(5)/D(4)$	1.1720	.2341	−.4272

Regression equation: $q(x)/D(j) = A(i) + B(i) \cdot (P[1]/P[2]) + C(i) \cdot (P[2]/P[3])$

APPENDIX TABLE 2 Regression coefficients to be used in estimating t^* for the case of declining mortality, where t^* is the number of years prior to the survey to which each mortality estimate corresponds and children ever born are classified by age of mother

Age group	$P(j)/P(m)$	$A(i)$	$B(i)$
20–24	$P(1)/P(2)$	1.3999	5.9156
25–29	$P(2)/P(3)$	1.1637	6.4668
30–34	$P(3)/P(4)$	−0.4262	10.1371

Regression equation: $t^*(i) = A(i) + B(i) \cdot (P[j]/P[m])$

APPENDIX TABLE 3 Regression coefficients (for North mortality pattern) to be used when estimating multipliers for *true* cohorts where children ever born are classified by age of mother

Age group	$q(x)/D(j)$	$P(m)/P(j)$	$A(i)$	$B(i)$
20–24	$q(2)/D(2)$	$P(1)/P(2)$	1.1635	−1.0530
25–29	$q(3)/D(3)$	$P(2)/P(3)$	1.1833	−0.4924
30–34	$q(5)/D(4)$	$P(3)/P(4)$	1.3408	−0.5210

Regression equation: $q(x)/D(j) = A(i) + B(i) : (P[m]/P[j])$

SOURCE: National Academy of Sciences, Committee on Population and Demography (1977).

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