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Relationship Between Infant Mortality and Socioeconomic Factors in Urban Areas

By Marion E. Altenderfer and Beatrice Crowther

This is the second of a series of studies dealing with the relationship between mortality and socioeconomic factors in the cities and counties of the United States. The study was prepared by the Division of Public Health Methods, Public Health Service.

The reduction in infant mortality is one of the outstanding public health achievements of the twentieth century. In 1915 the number of infant deaths per 1,000 live births was 99.9 in the birth registration area (10 States and the District of Columbia). By 1933, the first year for which all the States were included in the registration area, the rate had declined to 58.1. The downward trend has continued; the rate was 47.0 for 1940 and 33.8 for 1946.

This record low infant mortality rate for the United States as a whole conceals great variations among the sections of the country and among different groups of people. In 1946, 9 States had rates of 40.0 or more for all races combined, including 1 State with a rate of 78.2. The nonwhite rate for the United States was 49.5, and in 5 States more than 100 out of every 1,000 nonwhite infants died before they reached 1 year of age. Similarly, State rates conceal variations from county to county and city to city. These variations are associated with differences of many kinds among the various localities. The first paper of this series showed that in the cities of 100,000 or more population infant mortality varied inversely with per capita income and with the percentage of white persons in the population *(1)*.

The present study is an attempt to investigate the type of association between certain socioeconomic factors and infant mortality. It is based upon the smallest, most homogeneous units for which infant mortality data are available. These units are the urban places of 10,000 or more population; they will be referred to as cities, although some are called towns or townships.

Of the various measurable socioeconomic factors, four were chosen for analysis. Others were considered, but had to be discarded be-cause the data needed were not available for the cities in this study. The factors are: racial composition, income level of the population, size of community, and utilization of medical facilities. Since the components of the nonwhite population other than Negro are negligible in most cities, the percentage of white persons is used to measure racial composition (*2*). The per capita income is used as the measure of the relative income level of the population (*3*). The actual 1940 population is used as the measure of size of community (*4*). The percentage of births to residents which were attended by a physician in the hospital is used as the measure of the availability and utilization of medical facilities (*5*).

All the socioeconomic data are for the census year 1940, since three of the four factors are related to population. The infant mortality rates are for the 2-year period 1939–40 (*6*). If data had been available, a longer period (3 or even 5 years) would have been desirable to minimize further the year-to-year fluctuations in the number of infant deaths.

Since both the percentage of births hospitalized and the infant mortality rate are based on the number of live births registered, these factors are affected by differences in the completeness of birth registration. A birth registration study made in connection with the 1940 census showed that for the urban places of 10,000 or more population registration was 97 percent complete. Correction for incompleteness of birth registration would reduce the percentages of births hospitalized and the infant mortality rates very little for the 70 percent of the cities in which 95 percent or more of the births were registered. These factors would be reduced appreciably only in the 5 percent of the cities in which less than 80 percent of the births were registered. It is felt, therefore, that correction of the data for incompleteness of birth registration would have little effect on either the direction or the magnitude of the associations noted in the present study.

The relationship between infant mortality and the four socioeconomic factors is shown in the table. The 973 cities [1] included in the study were ranked according to their infant mortality rates and divided into quartile intervals. The three lower quartile intervals each contain 243 cities and the highest interval 244. For each quartile interval the rates measuring percentage of white persons in the population, per capita income, and percentage of births hospitalized were averaged. For the size-of-community factor the median rather than the mean population was used to avoid having the results unduly weighted by the large cities.

An inverse or negative association between infant mortality and per capita income, percentage of births hospitalized, and percentage of white persons in the population is evident. That is, a low infant

[1] Cities for which income data were not available are excluded.

Relationship between infant mortality and four socioeconomic factors, 973 cities

Quartile interval by rank of infant mortality rate	Mean				Median size of city, 1940
	Infant deaths per 1,000 live births, 1939–40	Per capita income, 1940	Percentage of births hospitalized, 1940	Percentage of white persons in population, 1940	
Total	47.4	$682	73.4	92.1	20,546
Lowest	27.8	722	86.5	97.7	21,954
Second	39.4	727	78.9	95.3	23,642
Third	49.0	686	72.4	92.2	22,474
Highest	73.2	595	55.8	83.2	16,787

mortality rate is associated with high values of these three socioeconomic factors and a high infant mortality rate with low values. The association with size of city is irregular; however, the median size of city for the highest quartile interval is considerably smaller than the other three medians.

The table also shows the relationship among the various socioeconomic factors. For example, a high percentage of births hospitalized is associated with a high per capita income and a high percentage of white persons in the population. For this reason the inverse association of some of the factors with infant mortality may be a result of direct association among the socioeconomic factors. To investigate this possibility, the relationship between infant mortality and each of the four factors must be studied with the effect of one or more of the other factors removed.

Standardized rates afford one way of eliminating individual factors.[2] The actual and standardized infant mortality rates for the 973 cities classified by each of the 4 socioeconomic factors in turn appear in the chart.[3]

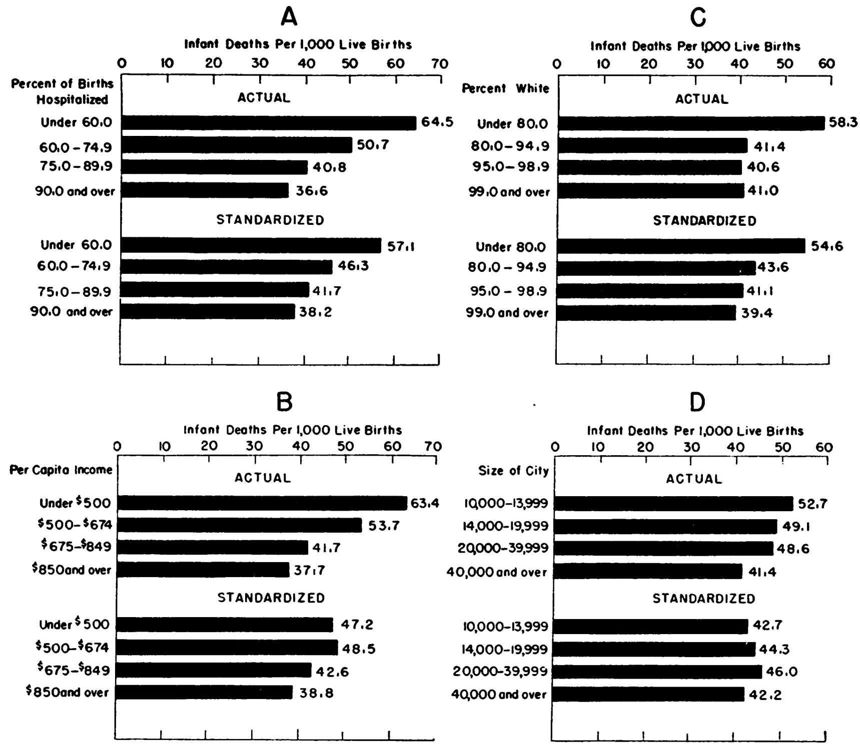
The percentage of births hospitalized shows a marked inverse association with infant mortality (chart, pt. A). When the other three socioeconomic factors are held constant, this relationship is still present. The relative differences in the rates between the class intervals are approximately the same as for the actual rates, but the range decreases from 27.9 to 18.9.

Elimination of the three other factors greatly reduces the negative association between per capita income and infant mortality (pt. B). Not only is the range of the standardized rates about one-third that of the actual rates, but the highest rate is now in the $500–674 income class, rather than in the lowest class.

Standardizing the rates classified by percentage of white persons in the population (pt. C) results in little change from the actual rates and has practically no effect on the range. Although the difference in

[2] This method is described in appendix I.

[3] The number of cities in each class interval for each factor appears in appendix II.



Actual and standardized infant mortality rates, 973 cities 1939–40.

the rates between the first 2 groups (those with less than 95.0 percent white population) has been reduced by standardization, it is still substantial.

Standardizing the rates by size of city (pt. D) reverses the trend in the actual rates. In other words, the slight negative association between size of city and infant mortality found in the actual rates has been replaced by an even slighter positive association.

The standardized rates thus show the association between infant mortality and each of four socioeconomic factors independent of the other three factors. The chart indicates that there is a greater inverse association with infant mortality for some of the factors than for others. Since this method of analysis does not permit a precise measure of the comparative importance of the various factors, correlation coefficients were used to evaluate each factor.

The simple or zero-order correlation coefficients between infant mortality and each of the four socioeconomic factors are:

Percentage of births hospitalized	−.57
Percentage of white persons	−.41
Per capita income	−.33
Size of city	−.057

These results are about what would be expected from an inspection of the actual rates in the chart. The correlation coefficients show that the degree of inverse association between infant mortality and percentage of births hospitalized is considerably greater than for the other three factors. The second highest order of association is with percentage of white persons in the population. The coefficient of correlation with per capita income is smaller but still large enough to have some meaning. The coefficient for size of city is not meaningful. The simple correlation coefficients measure the gross relationship between infant mortality and each of the four factors.

Partial correlation coefficients are used to measure the net relationship between infant mortality and each of the four factors with the effect of one or more of the other factors removed. Since the zero-order coefficient between infant mortality and size of city was not meaningful, neither the higher-order coefficients for this factor nor the effects of its removal on the other factors are shown here.[4]

The first-order partial correlation coefficients for infant mortality and percentage of births hospitalized are:

Effect of per capita income removed	—. 49
Effect of percentage of white persons removed	—. 47

Both are lower than the corresponding zero-order coefficient but still large enough to have meaning. Therefore the negative association between infant mortality and percentage of births hospitalized is largely independent of per capita income and of the color composition of the population.

The first-order coefficients for infant mortality and percentage of white persons in the population are:

Effect of per capita income removed	—. 35
Effect of percentage of births hospitalized removed	—. 20

The first of these coefficients is large enough to be meaningful, but the second is probably too small. In other words, the fairly high order of correlation between infant mortality and color composition (—.41) is reduced considerably when the effect of per capita income is removed and is cut almost in half when the effect of percentage of births hospitalized is removed.

The first-order correlation coefficients for infant mortality and per capita income are:

Effect of percentage of white persons removed	—. 26
Effect of percentage of births hospitalized removed	—. 055

Neither of these coefficients can be considered meaningful. The asso-

[4] See appendix III for complete set of the zero and higher-order correlation coefficients between infant mortality and the four socioeconomic factors.

ciation between infant mortality and per capita income is reduced to a negligible amount by removing the effect of percentage of births hospitalized.

The only second-order correlation coefficient large enough to have meaning (—.41) is that between infant mortality and percentage of births hospitalized with the effect of per capita income and percentage of white persons removed.

Summary

The association between infant mortality and four socioeconomic factors has been studied, using data for 973 cities of 10,000 or more population. Two methods are used to measure the association between infant mortality and each of the four factors with the effect of three factors removed.

The method of standardized rates shows that there is marked inverse association between infant mortality and percentage of births hospitalized when the combined effect of per capita income, percentage of white persons in the population, and size of city is removed. The negative associations between infant mortality and percentage of white persons in the population and between infant mortality and per capita income are less marked when the effect of the other three factors is removed. The inverse association with size of city disappears completely with standardization.

The method of partial correlation coefficients shows that the relationship between infant mortality and percentage of births hospitalized is independent of racial composition, per capita income, and size of community. The correlation between infant mortality and each of the other three factors becomes negligible when the effect of hospitalization is removed. In other words, the association of low infant mortality with high per capita income and with high percentage white is a consequence of the association of high values of these factors with high percentages of births hospitalized.

The high correlation between infant mortality and percentage of births hospitalized does not necessarily indicate a causal relationship. It is not hospitalization in the narrow sense of the word but the concomitant circumstances—such as an aseptic environment, the availability of skilled care and of facilities for dealing with emergencies, and the better prenatal care usually associated with hospitalization—that save many infant lives. The components of maternity care are difficult to evaluate individually but hospitalization at delivery is an index of the utilization of the best available care.

Infant mortality is affected by many factors in addition to those used in this analysis. Therefore, some of the association noted here may be due to factors that have not been considered. Nevertheless, hospitalization at delivery, with the high quality of infant and maternal care it provides, appears to be important in preventing infant deaths in this country.

REFERENCES

(1) Altenderfer, Marion E: Relationship between per capita income and mortality, in the cities of 100,000 or more population. Pub. Health Rep. 62: 1681–1691 (1947).

(2) Department of Commerce, Bureau of the Census: Sixteenth Census of the United States: 1940, Population, Vol. II. U. S. Government Printing Office, Washington, 1943.

(3) Survey of buying power, 1940. Sales Management 48: 84–283 (1941).

(4) Department of Commerce, Bureau of the Census: Sixteenth Census of the United States: 1940, Population, Vol. I. U. S. Government Printing Office, Washington, 1942.

(5) Department of Commerce, Bureau of the Census: Vital Statistics of the United States: 1940, Part II. U. S. Government Printing Office, Washington, 1943.

(6) Department of Commerce, Bureau of the Census: Vital Statistics of the United States, Supplement 1939–1940, Part III. U. S. Government Printing Office, Washington 1943.

Appendix I

Method of Deriving Standard or Adjusted Rates

This method is an adaptation of that known as "age adjustment by the direct method" used in mortality and morbidity statistics. A discussion of this method will be found in chapter IV of Vital Statistics Rates in the United States 1900–1940, by Forrest E. Linder and Robert D. Grove. The distribution of live births in the 973 cities was used as the "standard." The explanation below describes the steps taken in standardizing the rates for percentage of births hospitalized, holding per capita income, percentage white, and size of city constant (chart, part A).

1. The 973 cities were classified according to the four socioeconomic factors, using the four class intervals for each factor shown in the chart. This process resulted in 256 groups of cities. The infant mortality rates were computed for each group by dividing the total number of infant deaths by the total number of live births.

2. The 973 cities were then classified according to three factors, per capita income, percentage white, and size of city, using the same class intervals as above; 64 groups of cities were obtained. The live births were totaled for the cities in each of these groups.

3. Each of these 64 totals was multiplied in turn by the infant mortality rates for the four percentage-of-births-hospitalized groups

(from step 1) that fell into a corresponding classification by size of city, percentage white, and per capita income.

4. The results of the multiplications were totaled for each of the four percentage-of-births-hospitalized groups. Each of the totals was divided by the total number of live births in the same percentage-of-births-hospitalized group.

The four rates obtained are the infant mortality rates in each percentage-of-births-hospitalized group, adjusted for the other three factors. A similar procedure was used to obtain the standardized rates in parts B, C, and D of the chart.

Appendix II

Class Intervals

Several considerations influenced the choice of class intervals for the classification of the 973 cities according to the four socioeconomic factors. An attempt was made to have approximately the same number of cities in each of four groups and at the same time have the intervals begin and end at natural breaking points in the scale. This goal was achieved fairly well in the classification by size of city, the four groups containing 252, 221, 250, and 250 cities. For the classification by percentage of births hospitalized the groups are less uniform in size—235, 200, 283, and 255 cities. For the other two factors this method of classification could not be used. For per capita income the numbers of cities in the two end intervals were made approximately equal and the two middle intervals were made equal in span resulting in 122, 330, 394, and 127 cities in the respective groups. The large number of cities with 99.0 or more percent white persons in the population made impossible any systematic division into class intervals for racial composition; the numbers of cities in each group are 139, 205, 275, and 354.

Appendix III

Partial Correlation Coefficients—Infant Mortality and Four Socioeconomic Factors

1—infant mortality
2—percentage of births hospitalized
3—per capita income
4—percentage of white persons in the population
5—size of city

The subscript 12.3 indicates that the correlation is between infant mortality (1) and percentage of births hospitalized (2) with per capita income (3) held constant. The other subscripts have analogous meanings.

Zero order	First order	Second order	Third order
	$r_{12.3} = -.49$	$r_{12.34} = -.41$	
$r_{12} = -.57$	$r_{12.4} = -.47$	$r_{12.35} = -.49$	$r_{12.345} = -.41$
	$r_{12.5} = -.57$	$r_{12.45} = -.47$	
	$r_{13.2} = -.055$	$r_{13.24} = -.050$	
$r_{13} = -.33$	$r_{13.4} = -.26$	$r_{13.25} = -.055$	$r_{13.245} = -.048$
	$r_{13.5} = -.33$	$r_{13.45} = -.25$	
	$r_{14.2} = -.20$	$r_{14.23} = -.20$	
$r_{14} = -.41$	$r_{14.3} = -.35$	$r_{14.25} = -.20$	$r_{14.235} = -.20$
	$r_{14.5} = -.41$	$r_{14.35} = -.35$	
	$r_{15.2} = -.0061$	$r_{15.23} = -.00020$	
$r_{15} = -.057$	$r_{15.3} = -.012$	$r_{15.24} = -.020$	$r_{15.234} = -.015$
	$r_{15.4} = -.071$	$r_{15.34} = -.020$	

Effect of Sodium Fluoroacetate (1080) in Poisoned Rats on Plague Diagnosis Procedures

—Preliminary Report—

By I. Gratch, M. D., P. L. Purlia, and M. L. Martin, B. A.*

A preliminary report from experimental work indicates the rodenticide sodium fluoroacetate, 1080, does not interfere with the diagnosis of plague in poisoned rats. The suspicion of some workers that the presence of the poison in the body of a rat might interfere with plague diagnosis in the animal led to the present experiments. The Division of Foreign Quarantine of the Public Health Service initiated use of 1080 as a rodenticide on surface vessels in 1945.[1]

The experimental work concerning the following two problems was begun in March 1948, and is being continued:

1. When pathological material from rats poisoned with 1080 and suspected of plague infection is inoculated into guinea pigs, is there risk of killing the guinea pigs by means of 1080 rather than by *Pasteurella pestis*, thus making the diagnostic procedure unreliable?
2. Has 1080 any bactericidal or at least bacteriostatic properties, particularly against *P. pestis?*

*Surgeon (R), Inspection Officer, and Medical Technician, respectively, Public Health Service Quarantine Station, New York.

[1] John H. Hughes: Studies in deratization of surface vessels by means of 1080 (sodium fluoroacetate). Pub. Health Rep. 62: 933 (1947). Reprint 2793.

Of 1,190 rats autopsied during the present investigation, 202 were brought from ships treated with 1080; the others were collected after HCN gas fumigation, trapping, or manual killing, or were found dead from causes unknown to the writers. The 202 rats were delivered to the laboratory in 20 lots of from 1 to 35 rats; each lot was taken from a different ship. Liver and spleen from one rat, or from several rats, taken at random from most of these lots, served as material for the study in an attempt to answer the first of the two questions set forth above.

The technique of these experiments was as follows: Liver and spleen from one or more rats (presumably killed by 1080), mixed with a small amount of sterile physiologic solution of NaCl, were ground in a sterile mortar and allowed to sediment. Then the supernatant fluid was injected subcutaneously into one or more guinea pigs (0.5 to 1 cc. for each guinea pig) through the shaven skin of the abdominal region previously washed with soap and water and disinfected with alcohol.

None of the guinea pigs so inoculated presented any sign of illness during a period of observation which varied from 2 to 6 months.

It was concluded then that 1080 in rats killed by means of this poison does not reach the liver and spleen in sufficient quantity to kill a guinea pig by means of the above-described method. Therefore it was considered that, for the purpose of biological confirmation of tentative diagnosis of plague infection, routine inoculations of guinea pigs are justified not only in the case of rats killed by HCN gas fumigation or by trapping, but also in the case of rats poisoned with 1080.

Concerning the above observation, it could not be absolutely stated that the rats with which the experiment was conducted had ingested 1080. Was it possible that the rats had died from some cause other than poisoning by 1080?

Since, as far as our knowledge goes, there is no specific test which could unquestionably prove or disprove that the death of these rats was caused by ingestion of 1080, it was necessary to devise some modification in the procedure of the experiments. This was done in the following way: While liver and spleen of a rat supposedly dead from 1080 poisoning were injected into one guinea pig, the stomach of the same rat was used for injection into another guinea pig. For this purpose the stomach and its contents, mixed with a small amount of sterile saline solution, were ground in a sterile mortar and allowed to settle, and the supernatant fluid (0.5 to 1 cc.) was injected subcutaneously through the shaven skin of the abdominal region previously washed with soap and water and disinfected with alcohol.

The results of this series of experiments confirmed the previously stated conclusions. While guinea pigs inoculated with liver and spleen emulsions survived without revealing any sign of illness during a long period of observation, other guinea pigs injected with stomach and stomach content emulsions were seized, in about one hour after injection, with violent convulsions lasting 10 to 15 minutes, followed by death. These reactions are characteristic of 1080 poisoning. This experiment has been repeated several times, always with the same result. For the purpose of control, one guinea pig was injected with the stomach of a trapped rat, and another guinea pig was injected with the stomach of a rat killed by HCN gas. Neither guinea pig died.

During the investigation this question presented itself: How much value would these inoculations have for the purpose of biological confirmation (in this particular case, of plague infection) if 1080, in concentration used for rat poisoning, should possess some bactericidal properties? If it possessed such properties, then even if the autopsied rat presented lesions macroscopically suggestive of plague infection it would not be possible to isolate plague bacilli from the guinea pig inoculated with viscera of the rat. This consideration introduces the second problem previously stated, namely, bactericidal or bacteriostatic properties of 1080.

Observations of growth, on ordinary media, of habitual intestinal flora from rats revealed identical results regardless of whether the rats had been killed by 1080 or by HCN gas. These observations suggested that 1080 probably had no germicidal properties. To verify this impression an experiment was set up by implanting strains of various organisms on ordinary media, and contemporarily on identical media to which had been added 1080 in the concentration used in poisoning rats. The growth both on ordinary media and on media to which 1080 was added was identical; no bacteriostatic property of 1080 was noted.

The experiment was varied by the addition of compound 1080 and substances of known germicidal action to suspensions of the same organisms in sterile physiologic saline solution. After varying periods of contact of these organisms with the substances, the suspensions were implanted on blood agar. After 24 hours' incubation the plates were observed for growth, and revealed no bacteriostatic effect from presence of 1080. Plates implanted with organisms exposed to the known germicidal agents showed no growth.

Analagous experiments were conducted to test for any bacteriostatic properties of 1080 against *P. pestis.* For this purpose a nonpathogenic strain (A 1122) of *P. pestis* was used. The results were: absence, at least *in vitro,* of any bacteriostatic property in 1080 against *P. pestis.*

A final experiment *in vivo*, aimed to corroborate our assertion that 1080 in concentration used for rat poisoning has no bacteriostatic property against *P. pestis*, was conducted in the following way: One guinea pig was inoculated, subcutaneously, with 0.5 cc. of 24-hour broth culture of a virulent strain (Shasta) of *P. pestis* on October 26, 1948. Two days later when the guinea pig showed signs of illness, which we suspected to be plague infection, 1080 in water solution (18 grams per gallon) was administered orally to the animal by pipette. In about 35 minutes the guinea pig was seized with violent convulsions (characteristic of 1080 poisoning) and died. Autopsy, performed immediately after death, revealed plague infection. The liver and spleen and the bubo present in the left inguinal region were ground in a mortar containing a small amount of sterile saline solution. The emulsion obtained was left to settle, and 0.5 cc. of the supernatant fluid was injected subcutaneously into another guinea pig. This second guinea pig died from plague 4 days later, presenting on autopsy typical signs of the disease, including buboes in inguinal and axillary regions, hemorrhagic spots in subcutaneous tissue and on abdominal and thoracic viscera, and congested liver and spleen. Smears taken from liver and spleen showed microscopically small gram negative bacilli, morphologically similar to *P. pestis*. Cultures were made from the blood, heart, liver, lung, spleen, and buboes, all of which gave growth to organisms, later identified as *P. pestis*.

Summary and Conclusions

1. Experiments showed that sodium fluoroacetate (1080) in rats killed by this poison does not reach the liver and spleen in a sufficient amount to contraindicate routine injections from the rats into guinea pigs for biological confirmation of a tentative diagnosis of plague.

2. Sodium fluoroacetate, in concentration used for rat poisoning, has no bacteriostatic properties against *Pasteurella pestis*.

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"Infection Unit" and "Index of Aggregation"

Suggested Epidemiological Terms

By Filip C. Forsbeck, M. D.*

It is useful to divide typhoid cases into endemic and epidemic groups in order to study their relative rate of decline. Or, the typhoid phorologist may wish to compare accomplishment in tracing sources of endemic and epidemic cases. To find the source carrier of a 50-case outbreak is a perfect epidemiological record, but to find only 1 source carrier in the investigation of 50 endemic cases is ordinarily considered poor work. Yet in each instance the carrier discovery rate is 2.0 per 100 cases.

An added difficulty arises from the fact that it is not always obvious as to whether a given group of cases constitutes an outbreak. The Michigan criterion for classification as an outbreak is reasonable evidence of an indirect mode of transmission, irrespective of the number of cases.

Prior to January 1, 1936, the Michigan Department of Health arbitrarily chose three cases as the minimal for classification as an outbreak. Emphasis was thus placed on the fact that ordinarily the more cases involved, the greater the significance of an epidemiologic diagnosis of an outbreak. G. W. Anderson (*1*) has cited illustrations of single cases almost certainly traced to a mode of infection such as milk, and this situation has also occurred. in Michigan. He insists that there is thus no essential difference in one case or many cases insofar as mode of infection is concerned. The writer agrees with Dr. Anderson's concept: namely, that an outbreak should refer to one or more cases with probably a common mode of infection. This of course would mean that a large percentage of typhoid cases now classified as endemic would be classified as outbreaks if it could be determined that infection had taken place otherwise than by direct contact. There is indirect evidence that many single cases are thus caused, and in the case of endemic typhoid fever the classification, "direct contact" is almost always untenable. It is believed that the great bulk of endemic cases should be classified as "mode unknown."

The establishment of such criterion does not eliminate all conjecture, however. Assume, for example, that a typhoid carrier visits a family on August 1 only. Every member then contracts typhoid fever with onset dates on August 5, 9, 10, 12, and 14. It is true that all 5 cases

From the Michigan Department of Health and the Public Health Service.

* Dr. Forsbeck, at the time of his death in 1940, was a Surgeon in the Public Health Service; he had previously served as Director of the Bureau of Epidemiology of the Michigan Department of Health. After his death, Mrs. Forsbeck entrusted his papers to Dr. Gaylord W. Anderson at the University of Minnesota in the hope that among them might be material suitable for publication. This article, based on Dr. Forsbeck's Michigan studies and bearing the stamp of approval of the Michigan Department, was among these papers. It is published here exactly as Dr. Forsbeck left it, except for minor editorial changes and a redrawing of the graph which had been left merely as a rough sketch.

may have been primary, perhaps infected by food prepared by the carrier. However, assuming a minimum incubation period of 4 days, a maximum period not less than 2 weeks, and no chance of infection from outside the family, transmission of the causative organism may have occurred in 31 other combinations.

It is suggested that such difficulties may be obviated by treating single cases, and groups of two or more however related cases, as an "infection unit." Each of the following would be an example of an infection unit:

1. Any single isolated case.
2. A family group of cases, irrespective of how infection took place (unless by chance some of the cases had an entirely separate outside source).
3. Two apparently unassociated cases 10 miles apart at the same time on an island where typhoid has not occurred for 20 years.
4. An outbreak, including secondary and succeeding cases, and the causative case, if any.

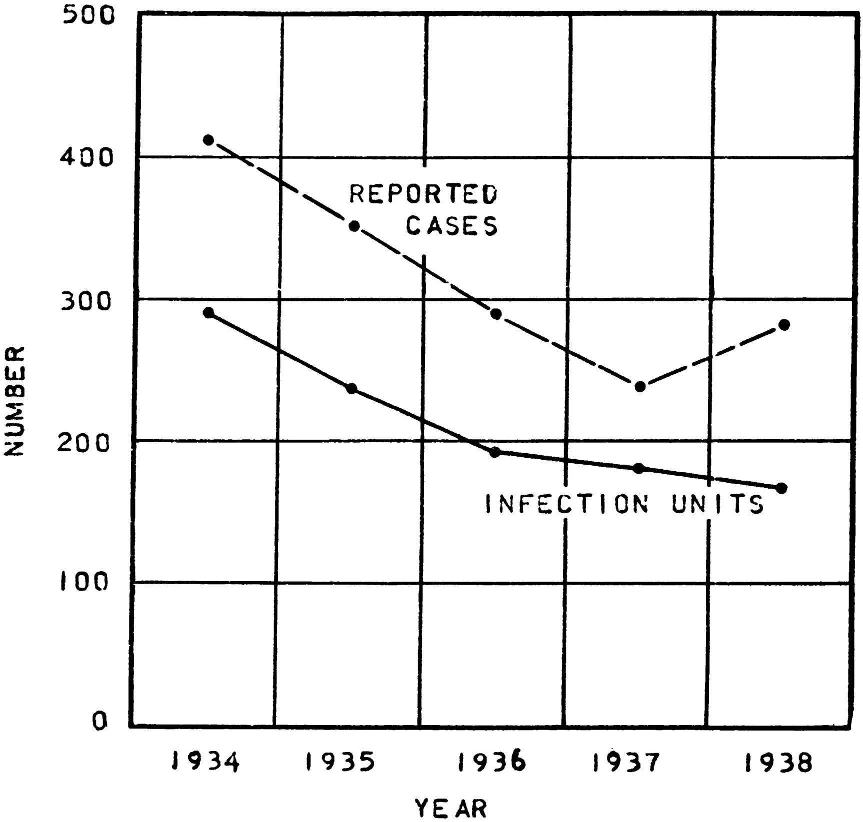
Unfortunately, we have been able to compute typhoid fever infection units in Michigan for only the years 1934–38. As may be noted in table 1, about 5 of every 6 infection units consist of single cases. The chart shows the sharp rise in case incidence from the trend line in 1938 while the infection unit incidence continues down in that year at about the same rate as in previous years.

Table 1. *Size of 1,063 typhoid infection units based on 1,571 typhoid cases reported in Michigan, 1934–38*

		Size of infection unit									
		1	2	3	4	5	6	7	8	9	10
Infection units	Number	888	94	29	19	8	4	5	4	2	3
	Percent distribution.	83.4	8.8	2.7	1.8	0.8	0.4	0.5	0.4	0.2	0.3
	Cumulative percent.	83.4	92.2	94.9	96.7	97.5	97.9	98.4	98.8	99.0	99.3
Cases	Number	888	188	87	76	40	24	35	32	18	30
	Percent distribution.	56.5	12.0	5.6	4.8	2.6	1.5	2.2	2.0	1.2	1.9
	Cumulative percent.	56.5	68.5	74.1	78.9	81.5	83.0	85.2	87.2	88.4	90.3

		Size of infection unit									
		11	12	13	14	15	16	17	23	49	Total
Infection units	Number	0	0	0	1	0	1	3	1	1	1,063
	Percent distribution.	0.0	0.0	0.0	0.1	0.0	0.1	0.3	0.1	0.1	100
	Cumulative percent.	99.3	99.3	99.3	99.4	99.4	99.5	99.8	99.9	100.0	------
Cases	Number	0	0	0	14	0	16	51	23	49	1,571
	Percent distribution.	0.0	0.0	0.0	0.9	0.0	1.0	3.2	1.5	3.1	100
	Cumulative percent.	90.3	90.3	90.3	91.2	91.2	92.2	95.4	96.9	100.0	------

Typhoid in Michigan, 1934–38, showing relative stability of infection unit trend compared with case incidence trend



	1934	*1935*	*1936*	*1937*	*1938*
Reported cases	411	351	288	239	282
Infection units	289	235	192	181	166
Aggregation index	1.42	1.49	1.50	1.32	1.44

The use of infection units in connection with carrier discovery rates is another illustration. It has been customary to measure this rate of accomplishment as total carriers, discovered by whatever means, per 100 reported cases.[1] By measuring the rate in terms of carriers discovered per 100 infection units, the misleading influence of a large outbreak is obviated.

For a number of years the rate at which carriers are discovered has been compared for Detroit, other full-time cities, county health units and part-time territory. These rates are compared with similar rates based on infection units in table 2. Although Detroit had the best record based on comparison of carriers discovered per 100 cases,

[1] In theory, a more logical measure would be the number of carriers found in connection with the investigation of cases, per 100 cases. In practice we have found that comparisons are not greatly affected by this procedure.

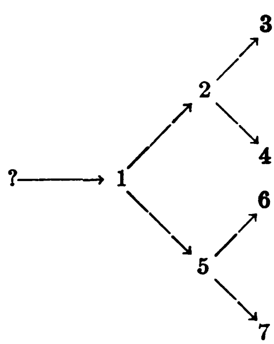
the full-time counties had the best record when measured by carriers discovered per 100 infection units. This shift is due to Detroit's remarkable freedom from outbreaks. Except for an outbreak of 60 cases in which infection was traced to another State and which is not included in the data, Detroit had but one outbreak in the period studied. For the period 1934–38, Detroit had but 1.16 cases per infection unit, other cities 1.49, counties 1.60, and part-time territory 1.54. The average number of cases per infection unit for Michigan, exclusive of Detroit, was 1.55 and for the entire State 1.48. It is suggested that this ratio of cases to infection units be termed the "index of aggregation".

Infection units and the index of aggregation may be of value as a unit of measure in the epidemiology and administrative practice of other diseases.

Table 2. *Typhoid cases reported, infection units, carriers discovered, and derived data by type of health jurisdiction, Michigan, 1934–1937*

Type of health jurisdiction	Cases reported	Infection units	Carriers discovered	Carrier discovery rate per 100 cases	Carrier discovery rate per 100 I. U.	Percentage deviation from mean	
						C.D.R. per 100 cases	C.D.R. per 100 I. U.
Detroit	189	161	33	17.4	20.5	36.0+	11.4+
All other full-time cities	214	148	15	7.0	10.1	45.3−	45.1−
County health units	322	204	53	16.4	25.9	28.0+	40.8+
Part-time territory	564	384	64	11.3	16.7	11.7−	9.2−
Michigan	1,289	897	165	12.8	18.4	--------	--------

Let us assume, for example, a series of seven cases of diagnosable tuberculosis in a small community and that infection actually took place as follows:



If investigation of the primary case had led to the discovery of all associated cases we would have 1 infection unit with 7 cases, and the index of aggregation would be 7.0. If, however, all associated cases had been discovered, but cases 3 and 4 independently of each other and other cases, we would have 7 cases, 3 units, and an index of aggregation of 2.3. Again, possibly only cases 1, 2, and 5 would be discovered and associated. The index of aggregation would then be 3.0. Indices of aggregation[2] for all possibilities are as follows:

Cases	Infection units						
	1	2	3	4	5	6	7
1	1. 0						
2	2. 0	1. 0					
3	3. 0	1. 5	1. 0				
4	4. 0	2. 0	1. 3	1. 0			
5	5. 0	2. 5	1. 7	1. 2	1. 0		
6	6. 0	3. 0	2. 0	1. 5	1. 2	1. 0	
7	7. 0	3. 5	2. 3	1. 7	1. 4	1. 2	1. 0

The index of aggregation might be a better measure of tuberculosis case finding than the ratio of new cases to deaths. Only experience would determine what constituted a satisfactory index. Certainly an aggregation index of 1.0 in a large city would indicate poor work in case finding.

Infection units and the index of aggregation may be useful in the study of outbreaks. In two outbreaks of dysentery for example, an index of aggregation of 1.1 would point less to contact infection than an index of 2.1. Frost (*2*) pointed out the importance of the secondary attack rate in relationship to the general attack rate. The infection unit of course includes the primary case in the household, but the size of the unit would be closely correlated with the secondary attack rate.

The number of infection units is readily determined by assigning an infection unit number to the first or index case of each single or associated group of cases.

Summary

The term "infection unit" is submitted as a useful epidemiological term to define any isolated case and any group of two or more associated cases, regardless of the manner in which they are associated.

The term "index of aggregation" is submitted as a useful epidemiological term to define the relationship of the number of cases to the number of infection units.

REFERENCES

(*1*) Anderson, Gaylord W.: Personal communications.

(*2*) Frost, W. H.: The familial aggregation of infectious diseases. Am. J. Pub. Health **28**: 7 (1938).

[2] In the example cited there is really only one index of aggregation, 7.0. Any other figure arrived at is an observed index of aggregation, not the true one.

INCIDENCE OF DISEASE

No health department, State or local, can effectively prevent or control disease without knowledge of when, where, and under what conditions cases are occurring

UNITED STATES

REPORTS FROM STATES FOR WEEK ENDED FEBRUARY 26, 1949

Of the total of 24,133 cases of measles reported for the week (last week 19,755, 5-year median 15,725), 16,522 cases (68 percent) occurred in the New England, Middle Atlantic, South Atlantic, and West South Central areas. Except in the New England area, increases, slight for the most part, were reported in all geographic divisions. The largest increases occurred in the Middle Atlantic (from 3,847 last week to 4,910), South Atlantic (2,950 to 4,139), and the West South Central (3,428 to 4,598). States reporting currently more than 728 cases are as follows (last week's figures in parentheses): Massachusetts 1,224 (1,396), New York 1,620 (1,482), Pennsylvania 2,562 (1,761), Wisconsin 818 (789), Kansas 902 (625), Maryland 1,361 (1,130), Virginia 1,388 (507), Texas 3,681 (2,684), California 1,100 (709). The total for the year to date is 133,503, same period last year 92,173, 5-year (1944–48) median 69,199.

A slight decline was recorded in the reported incidence of influenza, accounted for chiefly in the reports of South Carolina, 493 cases (last week 920), and Texas, 2,172 (last week 2,330). No other State reported more than 203 cases except Virginia, 470 (last week 329), and Arkansas, 292 (last week 235).

Of 57 cases of poliomyelitis, only 4 States reported more than 2 cases—California 14, Michigan and Texas 5 each, and Minnesota 3.

Two cases of anthrax were reported, in New York, and 2 cases of Rocky Mountain spotted fever, in Iowa.

For the year to date 10 cases of Rocky Mountain spotted fever have been reported (5-year median 4), and 257 cases of tularemia (5-year median 175).

Deaths recorded during the week in 94 large cities in the United States totaled 9,419 (last week 9,819), 9,811 and 10,200, respectively, for the corresponding weeks of 1948 and 1947, and a 3-year (1946–48) median of 10,200. The total for the year to date is 79,247, corresponding period last year 83,702. Infant deaths during the week totaled 649, last week 609, 3-year median 626. The cumulative figure is 5,408, last year 5,725.

Telegraphic case reports from State health officers for week ended February 26, 1949

(Leaders indicate that no cases were reported)

Division and State	Diphtheria	Encephalitis, infectious	Influenza	Measles	Meningitis, meningococcal	Pneumonia	Poliomyelitis	Rocky Mountain spotted fever	Scarlet fever	Smallpox	Tularemia	Typhoid and paratyphoid fever [d]	Whooping cough	Rabies in animals
NEW ENGLAND														
Maine	1	------	1	495	------	25	------	------	35	------	------	------	4	------
New Hampshire	------	------	2	8	------	4	------	------	4	------	------	------	4	------
Vermont	------	------	------	546	------	------	------	------	9	------	------	1	5	------
Massachusetts	9	------	------	1,224	1	------	------	------	300	------	------	------	52	------
Rhode Island	------	------	------	313	1	4	------	------	10	------	------	------	3	------
Connecticut	1	------	11	289	------	63	1	------	49	------	------	------	6	------
MIDDLE ATLANTIC														
New York	4	1	[b] 2	1,620	8	302	2	------	[c] 243	------	------	2	119	7
New Jersey	2	------	5	728	2	106	------	------	174	------	------	1	42	1
Pennsylvania	4	1	([b])	2,562	13	------	1	------	278	------	------	3	53	1
EAST NORTH CENTRAL														
Ohio	9	------	2	144	------	57	2	------	390	------	------	------	65	10
Indiana	8	------	4	130	2	60	2	------	133	------	1	------	42	17
Illinois	4	------	19	84	6	114	1	------	224	------	------	1	28	1
Michigan [a]	3	------	1	462	4	50	5	------	318	------	------	------	42	------
Wisconsin	------	------	73	818	4	10	1	------	87	------	------	------	30	------
WEST NORTH CENTRAL														
Minnesota	8	------	------	89	1	4	3	------	63	------	------	------	1	------
Iowa	------	------	------	30	------	4	1	2	38	------	------	------	4	12
Missouri	2	------	11	477	7	58	1	------	34	------	------	1	5	------
North Dakota	------	------	3	56	1	------	------	------	17	------	------	------	3	------
South Dakota	------	------	------	10	------	2	------	------	3	------	------	------	3	------
Nebraska	------	------	14	62	------	4	------	------	24	------	------	------	3	------
Kansas	1	------	15	902	------	9	2	------	56	------	1	------	2	1
SOUTH ATLANTIC														
Delaware	1	------	------	18	1	------	------	------	8	------	------	------	1	------
Maryland [a]	------	------	4	1,361	------	52	------	------	[c] 34	------	------	1	10	------
District of Columbia	1	------	------	76	------	20	------	------	4	------	------	------	2	------
Virginia	11	------	470	1,388	1	259	1	------	36	------	------	1	82	2
West Virginia	3	------	186	157	------	16	1	------	23	------	------	------	5	------
North Carolina	3	------	------	557	2	------	2	------	18	------	2	1	13	------
South Carolina	5	1	493	85	------	212	1	------	1	------	1	1	19	2
Georgia	4	------	16	378	3	90	2	------	13	------	3	4	7	8
Florida	8	------	6	119	1	13	1	------	6	------	------	1	2	3

See footnotes at end of table

Telegraphic case reports from State health officers for week ended February 26, 1949—Continued

Division and State	Diphtheria	Encephalitis, infectious	Influenza	Measles	Meningitis, meningococcal	Pneumonia	Poliomyelitis	Rocky Mountain spotted fever	Scarlet fever	Smallpox	Tularemia	Typhoid and paratyphoid fever [d]	Whooping cough	Rabies in animals
EAST SOUTH CENTRAL														
Kentucky	5		70	433	2	43	2		67			1	22	19
Tennessee	7		110	433	2	118			57		1	2	21	
Alabama	10		142	469	7	89			22		1		11	10
Mississippi [a]	1	1	18	83	1	22			2		3	2	9	
WEST SOUTH CENTRAL														
Arkansas	6		292	700	5	99			1		1	2	10	1
Louisiana	7		4	9	2	28	2		8		4	3	2	
Oklahoma	8		63	208	3	71			15		2	1	2	6
Texas	20		2,172	3,681	6	602	5		26			2	128	31
MOUNTAIN														
Montana	2	1	22	82					24		1			
Idaho			11	48		20	1		[c] 20				4	
Wyoming	3			24		11			8					
Colorado	1		11	312		59			19			2	3	1
New Mexico	6		2	303		32			20			1	2	
Arizona	3		203	162		53			16				17	1
Utah [a]			5	87		2	2		7				27	
Nevada	1													
PACIFIC														
Washington	1		10	347		19			49			1	1	
Oregon	2		18	464	1	56	1		23				13	
California	6	1	51	1,100	14	43	14		[c] 122			3	38	9
Total	181	6	4,542	24,133	101	2,905	57	2	3,138		21	38	967	
Median, 1944-48	261	8	6,425	15,725	175		26	1	3,288	8	20	41	2,251	
Years to date 8 weeks	1,491	59	36,750	133,503	708	19,368	760	10	22,145	8	257	323	8,436	
Median, 1944-48	2,443	61	92,191	69,199	1,643		314	4	24,382	58	175	333	17,994	
Seasonal low week ends	(27th) July 10		(30th) July 31	(35th) Sept. 4	(37th) Sept. 18		(11th) Mar. 20		(32d) Aug. 14	(35th) Sept. 4		(11th) Mar. 20	(39th) Oct. 2	
Since seasonal low week ends	6,605		73,020	185,896	1,552		28,087		44,843	23		3,442	18,469	
Median, 1943-48	10,009		135,749	95,323	3,147		13,690		62,953	134		4,571	43,221	

[a] Period ended earlier than Saturday. [b] New York City and Philadelphia only, respectively. [c] Including cases reported as streptococcal infections and septic sore throat.
[d] Including paratyphoid fever, currently reported separately, as follows: Vermont 1; Georgia 1; Colorado 1; California 1; salmonella infections, not included, were reported as follows: New York 2.

Anthrax: New York 2.
Deduction: Week ended Feb. 12, North Carolina, meningococcus meningitis, 1 case.
Additional: Week ended Feb. 19, measles, Arizona, 136 cases.
Alaska: Influenza 130; measles 4; pneumonia 4; streptococcal sore throat 6.
Territory of Hawaii: Diphtheria 1; influenza 655; measles 170; poliomyelitis 1; total of 602 delayed influenza reports for January and February included.

TERRITORIES AND POSSESSIONS

Panama Canal Zone

Notifiable diseases—December 1948.—During the month of December 1948, certain notifiable diseases were reported in the Panama Canal Zone and terminal cities as follows:

Disease	Residence [1]									
	Panama City		Colon		Canal Zone		Outside the Zone and terminal cities		Total	
	Cases	Deaths	Cases	Deaths	Cases	Deaths	Cases	Deaths	Cases	Deaths
Chickenpox	10		3		3				16	
Diphtheria	3								3	
Dysentery:										
Amebic	1	1	1				6	1	8	2
Bacillary	1		1		2		1		5	
German measles			1		1				2	
Hepatitis, infectious							1		1	
Malaria [2]	3		1		9		164		177	
Measles	2						1		3	
Meningitis, meningococcal	3	2	3			1	1		7	3
Mumps			1		2				3	
Pneumonia		9		3	16	3		8	[3] 16	23
Poliomyelitis			1		[4] 4	1	1		[4] 6	1
Tuberculosis		11		6	5			9	[3] 5	26
Typhoid fever							2	1	2	1
Typhus fever (murine)							1		1	
Yaws	1						3		4	
Yellow fever								5		5

[1] If place of infection is known, cases are so listed instead of by residence.
[2] 6 recurrent cases.
[3] Reported in the Canal Zone only.
[4] Includes 3 cases in which the infection was contracted in Philadelphia.

Puerto Rico

Notifiable diseases—4 weeks ended February 26, 1949.—During the 4 weeks ended February 26, 1949, cases of certain notifiable diseases were reported in Puerto Rico as follows:

Disease	Cases	Disease	Cases
Chickenpox	58	Syphilis	148
Diphtheria	45	Tetanus	9
Dysentery	2	Tetanus, infantile	1
Gonorrhea	177	Tuberculosis (all forms)	569
Influenza	144	Typhoid fever	12
Malaria	73	Typhus fever (murine)	2
Measles	80	Whooping cough	252
Poliomyelitis	1		

FOREIGN REPORTS

CANADA

Provinces—Communicable diseases—Week ended February 12, 1949.—During the week ended February 12, 1949, cases of certain communicable diseases were reported by the Dominion Bureau of Statistics of Canada as follows:

Disease	Prince Edward Island	Nova Scotia	New Brunswick	Quebec	Ontario	Manitoba	Saskatchewan	Alberta	British Columbia	Total
Chickenpox		19	21	255	939	39	73	78	332	1,756
Diphtheria		1	1	17	1	1		2		23
Dysentery, bacillary				1						1
German measles				113	27			5	6	151
Influenza		29			39		2		2	72
Measles		449	94	393	426	155	181	259	100	2,057
Meningitis, meningococcal					1				2	3
Mumps		28	4	118	470	42	54	19	66	801
Poliomyelitis				1						1
Scarlet fever		10	3	182	94	2	3	2	7	303
Tuberculosis (all forms)		3	7	80	16	6	16	20	38	186
Typhoid and paratyphoid fever				14	2		1			17
Undulant fever					1	1				2
Venereal diseases:										
Gonorrhea		3	12	87	53	26	10	30		221
Syphilis		3	5	77	36	12	8	7		148
Whooping cough		17		207	30	1	12	5		272

JAPAN

Notifiable diseases—5 weeks ended January 29, 1949.—During the 5 weeks ended January 29, 1949, certain notifiable diseases were reported in Japan as follows:

Disease	Cases	Deaths	Disease	Cases	Deaths
Diphtheria	1,882	232	Pneumonia	14,159	
Dysentery, unspecified	162	56	Scarlet fever	518	10
Gonorrhea	15,234		Smallpox	1	
Influenza	177		Syphilis	15,830	
Malaria	114	3	Tuberculosis	30,294	
Measles	6,199		Typhoid fever	528	49
Meningitis, epidemic	109	26	Typhus fever	32	
Paratyphoid fever	214	4	Whooping cough	5,453	

NOTE.—The above figures have been adjusted to include delayed and corrected reports.

POLIOMYELITIS

Ceylon.—Reports of poliomyelitis in Ceylon are as follows: Week ended November 27, 1948, 9 cases; period November 28–December 25, 1948, 21 cases.

Mauritius.—Poliomyelitis has been reported in the Colony of Mauritius as follows: Week ended January 15, 1949, 80 cases, 3 deaths; week ended January 22, 51 cases, 2 deaths.

REPORTS OF CHOLERA, PLAGUE, SMALLPOX, TYPHUS FEVER, AND YELLOW FEVER RECEIVED DURING THE CURRENT WEEK

NOTE.—Except in cases of unusual incidence, only those places are included which had not previously reported any of the above-mentioned diseases, except yellow fever, during recent months. All reports of yellow fever are published currently.

A table showing the accumulated figures for these diseases for the year to date is published in the PUBLIC HEALTH REPORTS for the last Friday in each month.

Cholera

India—Calcutta.—For the week ended February 12, 1949, 137 cases of cholera, with 56 deaths were reported in Calcutta, India, and during the week ended February 19, 106 cases, 48 deaths were reported.

India (French)—Pondicherry.—During the period January 30–February 12, 1949, 51 cases of cholera were reported in Pondicherry, French India.

Plague

India—Cawnpore.—During the week ended February 5, 1949, 14 cases of plague, with 4 deaths, were reported in Cawnpore, India.

Siam.—For the week ended February 5, 1949, 22 cases of plague, with 8 deaths, were reported in Siam.

Union of South Africa.—During the week ended February 5, 1949, 3 fatal suspected cases of plague were reported in the Union of South Africa, 2 cases at Uitenhage, about 20 miles from Port Elizabeth, and 1 case at Queensland, inland about 200 miles from Port Elizabeth.

Smallpox

Colombia.—For the period January 1–31, 1949, 227 cases of smallpox were reported in Colombia, including 15 cases in Bogota, 12 cases in Cali, and 11 cases in Medellin.

China—Amoy.—During the period January 21–31, 1949, 30 cases of smallpox were reported in Amoy, China, and during the period February 1–10, 32 cases were reported.

Ecuador.—For the period January 1–31, 1949, 215 cases of smallpox (including alastrim), with 17 deaths, were reported in Ecuador, of which 18 cases (alastrim) were stated to have occurred in Guayaquil and 14 in Quito.

India—Ahmedabad.—During the week ended February 5, 1949, 146 cases of smallpox, with 98 deaths, were reported in Ahmedabad, India.

Netherlands Indies—Sumatra—Palembang.—For the period January 23–February 19, 1949, 25 cases of smallpox were reported in Palembang, Sumatra, in the Netherlands Indies.

Typhus Fever

Brazil—Porto Alegre.—Correction: Reports of cases of typhus fever in Porto Alegre, Brazil, published in PUBLIC HEALTH REPORTS, November 12, 1948, p. 1,504, and January 14, 1949, p. 68, were in error. Later information states that these cases were typhoid fever.

Colombia.—During the period January 1–31, 1949, 256 cases of typhus fever were reported in Colombia, including 34 cases in Medellin.

Germany—Correction: The reports of murine typhus fever in the Bremen area (Land Bremen) in the United States Zone of Germany (see PUBLIC HEALTH REPORTS, January 7, 1949, p. 38, January 21, 1949, p. 92, and February 11, 1949, p. 200) were in error. Information dated February 16, 1949, states that these cases were typhoid fever.

Palestine.—Information from Jerusalem dated February 21, 1949, states that an outbreak of typhus fever is occurring in Palestine in the Hebron and Bethlehem districts, where about 100 cases had been reported since the middle of January.

Yellow Fever

No reports of yellow fever were received during the current week.

DEATHS DURING WEEK ENDED ABOUT FEB. 19, 1949

[From the Weekly Mortality Index, issued by the National Office of Vital Statistics]

	Week ended Feb. 19, 1949	Corresponding week, 1948
Data for 94 large cities of the United States:		
Total deaths	9,819	10,688
Median for 3 prior years	9,741	
Total deaths, first 7 weeks of year	69,828	73,891
Deaths under 1 year of age	609	776
Median for 3 prior years	776	
Deaths under 1 year of age, first 7 weeks of year	4,760	5,121
Data from industrial insurance companies:		
Policies in force	70,616,279	66,865,709
Number of death claims	13,167	14,490
Death claims per 1,000 policies in force, annual rate	9.7	11.3
Death claims per 1,000 policies, first 7 weeks of year, annual rate	9.8	10.8

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