STUDY OF CANCER MORTALITY IN ENGLAND AND WALES USING BIRTH-STANDARDIZED POPULATIONS

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It is well recognized that, in estimating secular trends of cancer mortality, the changing age-structure of the population must be taken into account. In England and Wales the proportion of the population in the older age groups has been steadily increasing; for example, in 1956-59, 53 per cent. of the female population were over 35 years of age, as compared with 41 per cent. in 1921-25. This increase is attributable to two causes: first, the falling mortality in lower age groups consequent upon improvements in public health and the treatment of infectious diseases; secondly, the decline in the birth-rate which took place between 1901 and 1931.

To allow for these changes in age-structure, it is customary to calculate mortality indices on a standard population which may be either hypothetical, or the population as it actually existed in a given year. Thus the standardized death rates published by the Registrar General were obtained by applying the age-specific rates in the current year to the 1901 population. The Comparative Mortality Index, which was introduced by the Registrar General in 1941, is the ratio of the rate in the current year to that in 1938 when both have been standardized on the mean of the proportionate age-distributions of the populations living in the two years. Both these measures of mortality therefore aim to present secular changes with the effects of changing age-structure eliminated. Thus, over the past decades, it has been shown that cancer mortality has been decreasing in women, but increasing in men, although male cancer mortality has also been decreasing if deaths from cancer of the lung are omitted. These trends in the mortality indices have been paralleled by the life experience of successive cohorts as demonstrated by Case (1956).

In considering mortality rates, it may be of considerable interest and importance to separate out

the two influences on age-structure mentioned above. For example, we may ask a different question from the one which standardized death rates and comparative mortality indices are designed to answer. We may wish to know whether the increased age of the population due to *decreased mortality in younger age groups* has in any way been compensated for by the decreasing cancer mortality mentioned above, and in particular, whether the chances that an individual *ultimately* dies from cancer have been increased by improvements in public health and the successes of medical science in treating infectious and other diseases.

One method of answering this question would be to construct life-tables for successive cohorts showing the number of deaths from cancer at each period of life. Thus the ratio, R , of the total deaths from cancer to the total number commencing in the cohort is the probability of ultimately dying from cancer. To perform this calculation on any one cohort, the age-specific mortality rates from cancer and all causes of death are required over a period of about 80 years, and reliable data over such an extended period is not as yet available. It is probable, for example, that the recorded deaths from cancer in England and Wales need considerable correction before 1900 (Cherry, 1925).

Errors in the diagnosis recorded on death certificates still occur (see Wynne-Griffith and Morgan, 1961 for a report of a recent investigation), but the overall cancer mortality rates since 1921 are not likely to be so much in error that useful information may not be obtained from studying their trends. A comprehensive set of age-specific mortality rates has been calculated by Case and Pearson for the period 1911 to 1955 and published in "Cancer Statistics for England and Wales, 1901-1955" (1957), but these do not enable any one cohort to be followed completely

from birth to death. Even if we restrict our inquiry to asking what is the probability of an individual's ultimately dying from cancer having reached a certain age (say 40), we can still only calculate this probability for one cohort. An alternative approach therefore seems desirable. The method of analysis suggested in this paper, while not able to provide the exact information that cohort analysis will eventually make available, does enable us to obtain at a much earlier date an approximate answer to the question that has been propounded.

CONSTRUCTION OF BIRTH-STANDARDIZED POPULATIONS

Fig. 1 shows the number of live births per year in England and Wales averaged over 5-year periods since 1841. The marked decrease between 1901 and 1931 is clearly seen. If we now consider the age-structure of the population in say 1921, then this must be closely related to the pattern of births per year during the preceding 90 years. For example, the persons aged 55–64 in 1921 are the survivors of those born in the 10-year period about 1861, the fraction surviving depending on the intervening life experience, *i.e.* the death rates at ages less than 55 and the effects of migration. If twice as many births had occurred around 1861, then assuming them subject to the same life experience, there would have been twice as many persons in the 55–64 age group in 1921. We may therefore construct a hypothetical

population that would have been alive in 1921 if equal numbers of births had occurred in each preceding year. To do this the number in a certain age group in the actual 1921 population must be multiplied by a weighting factor which is the inverse of the number of births per year that occurred around the period when the persons in that particular age-group were born.

Weighting factors for use with quinary-quinquennial age-groupings have been calculated from the births during the appropriate 9 years that form a quinary-quinquennial group (see Case (1956) for diagram showing how the groups are formed), giving weight to each year in proportion to the amount it contributes towards the formation of the group. As registration of births did not begin until 1837, the factor for 1836 was obtained by linear extrapolation. These factors were then applied directly to the mid-year populations given in "Cancer Statistics for England and Wales, 1901–1955" (1957). Thus the 1921 factor was used to multiply the numbers in the 0–4 age group in 1921–25, the 5–9 age group in 1926–30, the 10–14 age group in 1931–36, and so on. Calculations have been extended to cover also the 4-year period 1956–59, figures for 1960 not being available at the time when this work was carried out. The hypothetical population so obtained will for convenience be referred to as a "birth-standardized" population. The composition of the birth-standardized populations from 1921–59 is shown in the Table (opposite).

FIG. 1.—Live births per year in England and Wales, averaged over 5-year periods,

TABLE

AGGREGATE OF MID-YEAR BIRTH-STANDARDIZED POPULATIONS
 Figures in thousands; Standardized to a Live Birth Rate of 100,000 per year of each sex.

Sex	Age	Quinquennium							
		1921-25	1926-30	1931-35	1936-40	1941-45	1946-50	1951-55	1956-59
Male	0-	2,225	2,290	2,320	2,320	2,320	2,340	2,440	1,918
	5-	2,060	2,150	2,220	2,278	2,281	2,285	2,300	1,945
	10-	1,990	2,045	2,140	2,205	2,260	2,278	2,278	1,818
	15-	1,870	1,938	2,005	2,121	2,200	2,280	2,162	1,750
	20-	1,610	1,778	1,865	2,005	2,080	2,218	2,161	1,775
	25-	1,440	1,570	1,770	1,885	2,000	2,085	2,180	1,822
	30-	1,430	1,460	1,632	1,775	1,890	1,985	2,022	1,750
	35-	1,380	1,385	1,405	1,610	1,780	1,885	1,905	1,590
	40-	1,370	1,350	1,352	1,378	1,600	1,745	1,810	1,518
	45-	1,312	1,300	1,300	1,335	1,363	1,530	1,690	1,428
	50-	1,288	1,246	1,235	1,255	1,295	1,310	1,470	1,292
	55-	1,050	1,152	1,135	1,150	1,170	1,180	1,210	1,085
	60-	898	930	1,024	1,012	1,022	1,030	1,045	855
	65-	696	748	790	866	847	860	860	698
	70-	497	528	563	597	660	662	653	526
	75-	288	302	327	347	383	440	430	344
	80-	137	147	147	159	176	204	224	177
	85 and Over	54	53	57	61	76	87	89	86
Female	0-	2,280	2,340	2,365	2,342	2,360	2,358	2,462	1,931
	5-	2,121	2,200	2,280	2,338	2,330	2,338	2,330	1,969
	10-	2,058	2,115	2,205	2,280	2,338	2,318	2,320	1,838
	15-	1,990	2,028	2,098	2,218	2,280	2,340	2,342	1,840
	20-	1,870	1,960	2,030	2,119	2,218	2,290	2,375	1,900
	25-	1,822	1,838	1,960	2,035	2,110	2,200	2,296	1,909
	30-	1,752	1,770	1,814	1,945	2,010	2,098	2,160	1,842
	35-	1,660	1,690	1,715	1,784	1,915	1,988	2,058	1,721
	40-	1,625	1,616	1,650	1,690	1,758	1,870	1,940	1,640
	45-	1,480	1,555	1,570	1,620	1,658	1,720	1,835	1,528
	50-	1,398	1,410	1,478	1,520	1,570	1,592	1,670	1,431
	55-	1,194	1,285	1,138	1,400	1,454	1,474	1,528	1,280
	60-	1,052	1,088	1,172	1,212	1,305	1,330	1,392	1,149
	65-	870	936	980	1,046	1,082	1,156	1,215	1,009
	70-	682	716	762	814	863	910	990	825
	75-	453	470	501	541	585	636	686	596
	80-	238	256	260	290	319	355	386	338
	85 and Over	119	120	129	136	165	195	204	187

Fig. 2 (overleaf) shows how the birth-standardized population of England and Wales has aged between 1921 and 1955, by comparison with the actual population. It may be seen that over the period considered the proportions of the actual population over 35 and over 55 years of age have been increasing considerably, and this increase is maintained in the female birth-standardized population, but to a much smaller degree. In fact, about three-quarters of the increase in the actual population is attributable to the birth-rate variations. In males, however, the actual increase is seen to be less than for females, and the proportions of the birth-standardized populations over 35 and over 55 have remained almost constant.

We must consider at this point what other factors can influence the birth-standardized populations apart from changes in life-expectancy brought about by improvements in public health and medical treatment. The first factor is evidently the reliability of

the birth registration figures from which the weighting factors are calculated. Registration of births in England and Wales was not made compulsory until 1875, but Farr (1885) estimated that between 1837 and 1876 the proportion of unrecorded births amounted to only about 5 per cent., and that since 1876 the proportion had been much less. It is feasible that in the first years of registration the deficiency of the records was more than 5 per cent., although in 1842 the Registrar General in his 4th Report suggested that the numbers omitted amounted to only "several thousands", this in a total registration of about 500,000. The registration figures are not therefore likely to involve any error that will affect significantly the age-structure of the birth-standardized populations.

The second factor to be considered is the question of migration. This has varied considerably from decade to decade, giving rise, for example, to a net

FIG. 2.—Percentage of total population over 35 and over 55 years of age, averaged over 5-year periods.

loss of 60,100 persons per year in 1881–1891 and of only 6,900 per year in the following decade. These figures were 6·8 and 0·8 per cent. respectively of the live births per year during the same periods. Between 1931 and 1951 there was a net gain of 33,000 per year. In the “Royal Commission on Population Report” (1949) the effect of migration was considered in some detail, and the conclusion was reached that, within the limits of a net flow of not more than 100,000 persons per annum, the effect of migration on the age balance of the population “would be inconsiderable”. These limitations seem to have been complied with over the period considered in this paper.

The third factor is the loss in the male population caused by war. The 577,000 lost in the 1914–18 war would have been largely in the 25–44 age group in 1921–25 and therefore in the 55–74 age group in 1951–55. The 240,000 non-civilians who died outside England and Wales in 1939–47 would mostly have been under 55 in 1951. Assuming all the losses to be males, the approximate effect on the birth-standardized population may be calculated, and the curves corrected for these losses are also shown in Fig. 2. It is seen that the dip in the male over-35 curve and the constancy of the male over-55 curve are largely due to war losses. Without these the male birth-standardized population would have shown some ageing, but still not quite to the same extent as the female population.

In all that follows the birth-standardized populations have been used without correction for losses due to migration and war. In females, the secular trends in these populations may be taken to be a reasonable representation of the ageing that would have taken place because of the improvements that have been made in public health and the treatment of diseases in younger life. In males the ageing due to this cause is slightly underestimated.

CANCER MORTALITY RATES

A cancer mortality rate may now be calculated by applying the age-specific rates pertaining to a given period to the birth-standardized population constructed for that period. For convenience, the rate so obtained will be referred to as the Birth-Standardized Rate. Its values for the period 1921–59 are compared with the actual crude mortality rates in Fig. 3 (opposite), where cancer does not include Hodgkin’s disease or leukaemia. It may be seen that the birth-standardized rate for the female population has remained almost constant, *i.e.* the crude female cancer mortality rate for England and Wales has only increased during this period through variations in the number of live births per year from 1936 onwards. For the male population, the birth-standardized rate has increased, but the increase is more than accounted for by deaths from cancer of the lung. About two-thirds of the increase in the actual crude mortality rate has been due to birth-rate variations.

FIG. 3.—Death rate from all forms of cancer (excluding leukaemia and Hodgkin's disease), averaged over 5-year periods.

EFFECT OF BIRTH-RATE CHANGES ON CANCER MORTALITY RATES

It is of interest to consider what the cancer mortality rates over the past decades would have been if the pattern of births per year had been different from what it actually was. Supposing, for example, the annual number of female births had continued to increase after 1881 in accordance with the dotted line COD in Fig. 4, or alternatively that the births had been continually decreasing during the period before 1901 as shown by the line AOB in Fig. 4 (overleaf).

Populations and mortality rates corresponding to these assumptions have been calculated for 1921–59, in the same manner as the birth-standardized data, but using a different set of weighting factors. Fig. 5 (overleaf) shows that the fraction of the female population over 35 years of age would have increased with either continually falling or continually rising annual births, but that the increase would have been less than that which actually occurred. At first sight this result may be surprising, as it seems to contradict the common belief that a falling birth-rate would lead to the most severe ageing of the population. It must be

noted, however, that in any year the fraction over 35 years of age would have been considerably higher with a falling rate (AOB) than with a rising rate (COD), and that Fig. 5 shows only that the *change* in age-structure during the period 1921–59 would not have been so marked as it was in practice. It is in fact a sharp change in the trend of annual births that has the greatest influence on the variations in the number of people who are over 35 years old some 40 years after the change occurs. A continually decreasing number of annual births to about 1891 followed by a continually increasing number thereafter (line AOD in Fig. 5) would have led to a dip in the fraction who were over 35 years old around 1936–46, as shown in Fig. 5.

The crude female cancer mortality rates that would have arisen from these hypothetical patterns of annual births are shown in Fig. 6. It may be seen that the mortality rate arising from either a continually increasing or a continually decreasing number of births per year would have changed much less markedly than did the observed rate, and that, had we experienced a period of decreasing

FIG. 4.—Female live births per year in England and Wales, averaged over 5-year periods.

FIG. 5.—Percentage of female population over 35 years of age resulting from different patterns of birth-rate variations.

births followed by a period of increasing births, the Registrar General would have been able to report a progressively decreasing crude cancer mortality rate

over the last 35 years. This may well happen in about 20 years time if the upswing of the birth rate curve that began in 1936 is maintained.

FIG. 6.—Female cancer mortality rates resulting from different patterns of birth-rate variations.

PROBABILITY THAT AN INDIVIDUAL ULTIMATELY DIES FROM CANCER

We may now return to the question that we asked earlier, and inquire how the chances of ultimately dying from cancer have changed for the individual. In a birth-standardized population every cohort starts off with the same number of persons, and if the probability of dying from cancer in each cohort

remains constant, then the ratio, R' , of cancer deaths to total deaths in successive years will also remain constant. (The relationship between R and R' when R is changing will be discussed in the next section.) This ratio is plotted in Fig. 7 for the whole population and also for those over 35 years old. It may be seen that the ratio for both sexes of all ages has increased steadily over the period studied. For

FIG. 7.—Deaths from cancer as a fraction of deaths from all causes in the birth-standardized populations of England and Wales, averaged over 5-year periods.

females over 35 years of age, the ratio has remained almost constant during the past 20 years. This is of considerable interest in considering future trends, since the curve for all ages is evidently approaching that for the over 35s and may soon coalesce with it. Already in 1956-59, only 6.5 per cent. of the deaths from all causes and 2.3 per cent. of the deaths from cancer occurred in persons under 35 years of age in the female birth-standardized population. If deaths in persons under 35 had occurred only from cancer, the all-ages ratio would have been 0.177, as compared with 0.173 for the over-35s. It seems possible therefore that, if present trends continue, the probability of ultimately dying from cancer for the individual female may reach an equilibrium value of about 0.18, in spite of a progressive ageing of the population due to decreasing age-specific death rates at all ages. For females it will then no longer be true to say that "the ultimate chance of dying of some neoplastic disorder is increasing with our better standard of living" (Smithers, 1960). Two possibilities could effect this prediction. One is a marked advance in the treatment of some disease other than cancer that is at present responsible for a large proportion of deaths at ages over 35. So far it would appear that reductions in age-specific rates for female cancer have been of the same order as the reductions that have occurred in the rates for other diseases in persons over the age of 35.

The second possibility is that the trend of cancer death rates will be changed so that the female picture becomes similar to the male. It may be seen in Fig. 7 that the ratio for males over 35 has continued to increase over the whole period, this trend being entirely accounted for by deaths attributed to cancer of the lung. The age-specific rates for lung cancer have also been increasing in females (Case, 1956). Should lung cancer become a predominant cause of death in females, then the fall in age-specific rates for all types of cancer may be slowed and eventually reversed, and the probability that females ultimately die from cancer will then continue to increase. On the other hand, the curve for males over 35 years old in Fig. 7 shows some signs of becoming less steep, and may therefore also be approaching an equilibrium value in spite of the lung cancer "epidemic". The mortality figures for the next decade will be of particular interest, as they will show whether the constancy of the over-35 curve for females and the flattening of the over-35 curve for males are only temporary phenomena, or whether these trends are likely to be maintained. The joining up of the all-ages curves with the over-35s curves may also be achieved after another 10 years have elapsed.

RELATIONSHIP BETWEEN R AND R'

The relationship between the ratio, R , of cancer deaths to deaths from all causes in a cohort, and the corresponding ratio, R' , in the birth-standardized population at any time may be set out formally as follows:

Let N_0 = Number of live births per year, which is assumed constant;

N_t = Number of a cohort alive at an age t , i.e. N_t/N_0 is the proportion surviving to age t ;

T = Time in years measured from some arbitrary origin;

T_0 = Time at which R' is calculated.

The age-specific mortality rates from all causes of death will be a function of the age, t , for which the rate applies, and the time, T , at which the cohort was born. Let it be represented by a function $F(t, T)$.

$$\text{Then } -\frac{1}{N_t} \cdot \frac{dN_t}{dt} = F(t, T),$$

and by integration

$$N_t = N_0 \exp \left\{ - \int_0^t F(t, T) \cdot dt \right\}.$$

Let the age-specific mortality rate for deaths from cancer be represented by a function $f(t, T)$.

$$\therefore -\frac{1}{N_t} \cdot \frac{dN_c}{dt} = f(t, T), \text{ where } dN_c \text{ is the number}$$

of cancer deaths between the ages of t and $t+dt$.

The total deaths from cancer in a cohort born at time, T , is therefore

$$\begin{aligned} - \int_{t=0}^{t=\infty} dN_c &= \int_0^{\infty} N_t \cdot f(t, T) \cdot dt \\ &= N_0 \int_0^{\infty} f(t, T) \cdot \exp \left\{ - \int_0^t F(t, T) \cdot dt \right\} \cdot dt \\ \therefore R &= \int_0^{\infty} f(t, T) \cdot \exp \left\{ - \int_0^t F(t, T) \cdot dt \right\} \cdot dt. \end{aligned}$$

To obtain an expression for R' , consider first the deaths per year between the ages of t and $t+dt$ at time T_0 . These will occur in the cohort born at time $T_0 - t$, and we may therefore write:

$$\begin{aligned} -\frac{dN_t}{dt} &= N_t F(t, T_0 - t) \\ &= N_0 \cdot F(t, T_0 - t) \cdot \exp \left\{ - \int_0^t F(t, T_0 - t) \cdot dt \right\} \cdot dt. \end{aligned}$$

The total number of deaths per year at time T_0 is therefore given by:

$$\begin{aligned} \int_{t=0}^{\infty} dN_t &= N_0 \int_0^{\infty} F(t, T_0 - t) \cdot \exp \left\{ - \int_0^t F(t, T_0 - t) \cdot dt \right\} \cdot dt \\ &= N_0 \left[- \exp \left\{ - \int_0^t F(t, T_0 - t) \cdot dt \right\} \right]_0^{\infty} \\ &= N_0 \left[1 - \exp \left\{ - \int_0^{\infty} F(t, T_0 - t) \cdot dt \right\} \right]. \end{aligned}$$

Similarly, the deaths from cancer per year at time T_0

$$= N_0 \int_0^{\infty} f(t, T_0 - t) \cdot \exp \left\{ - \int_0^t F(t, T_0 - t) \cdot dt \right\} \cdot dt,$$

whence

$$R' = \frac{\int_0^{\infty} f(t, T_0 - t) \cdot \exp \left\{ - \int_0^t F(t, T_0 - t) \cdot dt \right\} \cdot dt}{1 - \exp \left\{ - \int_0^{\infty} F(t, T_0 - t) \cdot dt \right\}}.$$

We may consider three cases:

Case 1.—If $F(t, T)$ and $f(t, T)$ are independent of T , then R is a constant for all T , and

$$\exp \left\{ - \int_0^{\infty} F(t, T) \cdot dt \right\} = N_{\infty}/N_0 = 0 \text{ in any practical case.}$$

Hence $R' = R$, a result which may be seen intuitively.

Case 2.—If $F(t, T)$ is independent of T (equals $F(t)$ say), but the age-specific rates for cancer change from cohort to cohort, then again the denominator of R' is unity and

$$R' = \int_0^{\infty} f(t, T_0 - t) \cdot \exp \left\{ - \int_0^t F(t) \cdot dt \right\} \cdot dt.$$

This is not subject to any further analysis unless we assume some form for $f(t, T)$. Suppose $f(t, T) = (a + b \cdot T) \phi(t)$, i.e. the age specific rates at age t are a linear function of T .

Then

$$\begin{aligned} R &= \int_0^{\infty} (a + b \cdot T) \cdot \phi(t) \cdot \exp \left\{ - \int_0^t F(t) \cdot dt \right\} \cdot dt, \\ &= (a + b \cdot T) \cdot X, \end{aligned}$$

where

$$X = \int_0^{\infty} \phi(t) \cdot \exp \left\{ - \int_0^t F(t) \cdot dt \right\} \cdot dt,$$

and is independent of T .

$$\begin{aligned} R' &= \int_0^{\infty} \left\{ a + b(T_0 - t) \right\} \cdot \phi(t) \cdot \exp \left\{ - \int_0^t F(t) \cdot dt \right\} \cdot dt, \\ &= (a + bT_0) \cdot X - b \cdot Y, \end{aligned}$$

where

$$Y = \int_0^{\infty} t \cdot \phi(t) \cdot \exp \left\{ - \int_0^t F(t) \cdot dt \right\} \cdot dt;$$

whence we may see that $R' = R$ when $T_0 - T = Y/X$.

Thus R' and R are two identical linear functions of T , but with the origin of the T -axis displaced by an amount equal to Y/X .

X and Y have been evaluated using the computer in the University of Cambridge Mathematical Laboratory, on the assumption that $F(t)$ is a Gompertz function of the form $K \cdot e^{\lambda t}$, and similarly that $\phi(t)$ is of the form $e^{\lambda' t}$. The value chosen for λ was 0.092 years^{-1} , corresponding to a doubling time of 7.5 years, which is a reasonable approximation to the slope of the curves of log age-specific mortality for deaths from all causes against age for age groups above 55. It was found that the value of Y/X was comparatively insensitive to variations of both K and λ' . Thus for $\lambda' = 0.055 \text{ years}^{-1}$ (doubling time of 12.5 years), Y/X varied from 62.0 to 66.4 years as K varied from 7.5×10^{-5} to 5.0×10^{-5} . A reasonable value for K for the England and Wales female cohort born in 1871 would be about 6.4×10^{-5} . Keeping K constant at 6.0×10^{-5} , it was found that Y/X varied from 58.6 to 67.8 years as λ' increased from 0.035 to 0.065 years^{-1} . Thus, in this very idealized case, the value of R' would equal the value of R in a cohort born about 60 years earlier, whereas the complete cohort information would only be available on those born up to a date about 85 years earlier than the time at which R' is determined.

Case 3.—If $F(t, T)$ varies with T , while $f(t, T)$ remains constant for each cohort, then the expressions for R and R' cannot be further simplified in any useful manner, and an empirical approach has therefore been made. Fig. 8 (overleaf) shows the mortality curves for a series of hypothetical cohorts born at 10-yearly intervals. The curves for cohorts 1 to 13 approximate wholly or in part to the life experience of female cohorts in England and Wales from 1831 to 1951. Suppose we assume that $F(t, T)$ for Cohort 1 is the same for all cohorts born before Cohort 1, and that all those born after Cohort 15 have the same $F(t, T)$ as Cohort 15. We are postulating therefore a gradual change in $F(t, T)$ between $T=30$ and $T=170$ years on the scale chosen. By applying a constant series of cancer age-specific mortality rates to the cohorts as they progress through their life experience, we may calculate R for each cohort and R' throughout the period. The age-specific mortality rates used were approximately those for all forms of cancer (excluding leukaemia and Hodgkin's disease) in England and Wales during the quinquennium 1911–15.

The results of the calculations are plotted in Fig. 9 (overleaf). It may be seen that the value of R' follows that of R with a time-lag of about 40 years, narrowing to about 25 years as the upper bend of the curve is reached. The considerable advantage of the use of birth-standardized populations in a case similar to this is demonstrated by these curves. At $T=210$ years, the flattening out of the R' curve is apparent, but at this time complete information on

FIG. 8.—Hypothetical survival curves for a series of cohorts born at 10-year intervals.

only the cohorts up to, and including, Cohort 10 would have been available. Thus the curve of R would have only reached point A ($T=120$ years), and there would be no sign as yet of the reduction of the rate of increase of R , and the constant value that was eventually going to be reached.

This third case is the most nearly analogous to what has taken place in England and Wales during the last century. Death rates from all causes have fallen strikingly, particularly in the younger age groups, while the age-specific death rates from cancer have remained relatively constant. If the cancer death rates in older age groups now decrease in such a way as to keep pace with the fall in the death rates from all causes, then the rise in R would cease. Taking into account the time-lag, we might consider that the trend of R' in Fig. 7 for females shows that R may turn out to be approximately constant for the generations who have been born since 1931.

DISCUSSION

The object of this investigation was to see whether the mortality data for England and Wales lent any

support to the prevailing opinion that improvements in life expectancy, brought about by improvements in public health and medical treatment, are partly responsible for the increase in crude cancer mortality rates, and are providing the individual with an ever-increasing chance of ultimately dying from cancer. Smithers (1960) has already been quoted as an example of this opinion. Two other examples from United States sources are also worth quoting. Hammond (1959), in a paper discussing trends in cancer death rates and cure rates, shows that, in a number of countries including England and Wales, the crude cancer death rate has increased from 1930 to 1956. He then goes on to say that this increase is "due in large measure to the fact that the population . . . has become older, on the average, as a result of the improvement in life expectancy coupled with changes in birth rates". The other example is from a manual for students and physicians on human cancer by Black and Speer (1957), which opens with these remarks: "Each year, in the United States alone, more than 200,000 men, women, and children die of cancer. This figure may be expected to increase with

FIG. 9.—Deaths from cancer as a fraction of deaths from all causes.

the passage of time, owing to the increasing average life span and the decreasing number of deaths from infectious diseases". Such statements tend to be interpreted pessimistically by members of the general public, who are becoming increasingly concerned with the cancer problem.

The results of the present investigation show that the statistical data for England and Wales do not entirely support the views quoted above. In the first place, the increase in life-expectancy in females over the past 40 years cannot be held responsible for any of the increase in the crude mortality rates for female cancer. For males, only one-third of the increase in the crude cancer mortality rate is attributable to causes other than birth-rate variations, and this fraction is not so much due to increased life expectancy, as to the increase in the age-specific mortality rates for cancer of the lung. In the second place, it seems possible that, in spite of a continuing increase in life expectancy, the probability that an individual female ultimately dies of cancer will, if present

trends continue, reach an equilibrium value of about 1 in 5.5. Whether a similar possibility also holds good for males is not yet apparent, and in this connexion the mortality data for the next decade will be of particular interest.

SUMMARY

A method is described of constructing birth-standardized populations for England and Wales between 1921 and 1959. These populations have the age-structure that would have occurred in practice if the number of live births per year since 1836 had remained constant. Any ageing that occurs in the birth-standardized populations is therefore mainly due to improvements in public health and the treatment of disease in younger age groups, and the ratio of cancer deaths to deaths from all causes in these populations provides an indication of the probability that an individual ultimately dies of cancer. The cancer mortality calculated on these populations is

called the Birth-Standardized Rate, and it is shown that for females the rate has remained almost constant over the period studied. The rise in the rate for males is more than accounted for by deaths from cancer of the lung. It is also suggested that for females the probability of ultimately dying from cancer may be approaching an equilibrium value of about 0.18. The effect of birth rate variations upon the ageing of the population and upon crude cancer mortality rates is also discussed.

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