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Declining Old-Age Mortality

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ABSTRACT

This paper explores some policy implications of declining mortality among the elderly population. There are two views of how health progress will affect the social burden of caring for the aged. One holds that prolonging the lives of frail individuals will result in rapidly increasing medical and other costs per aged person. A second view suggests that health progress and behavioral changes will reduce both mortality and morbidity rates, lowering the average cost per person of caring for the aged. This paper investigates recent trends in the health status of the elderly to distinguish between these two views. The findings suggest that reductions in morbidity and mortality have been roughly counterbalanced by the rising frailty of the surviving population. Age specific institutionalization rates and medical cost patterns have been relatively stable for the last two decades, suggesting neither dramatic improvements nor sharp reductions in the health status of the elderly.

Recent years have witnessed dramatic improvements in the longevity of the elderly population. Life expectancy for women at age 65 increased by 3.3 years, a 22 percent change, between 1950 and 1980. Age-specific death rates fell by 29 percent for the over-85 female population, and by 19 percent for over-85 males. Continuation of these trends could have major implications for public policy towards the elderly. If mortality rates continue to drop, the elderly population will be substantially larger than if mortality rates remain constant at their current level. The number of extreme elderly, those over age 85, could rise especially rapidly.

This paper explores some policy implications of the dramatic longevity gains which have occurred, and which are likely to occur in the future. We focus on the potential burden, through demand for medical care and other resources, which the elderly are likely to create. Different views on this issue are possible. A pessimistic outlook would hold that improvements in life expectancy are likely to be associated with large increases in the costs of supporting the elderly. The elderly will grow more numerous particularly at very old ages where support costs are greatest. These costs may be particularly high for "marginal survivors" -- those who would have died at earlier ages but for recent progress in reducing mortality. A more optimistic view is also possible. It would argue that the same forces which have led to recent declines in mortality might also be expected to lead to declining morbidity and increased ability to function. Mortality reductions lower the number of the aged who are within a year or two of death. Since these are the years when support costs are highest,

especially for health care, the social burden per aged person might actually decline.

The relative importance of these two effects is an empirical question which cannot be resolved by a priori argument. Which view is more appropriate is, however, clearly an important policy issue. Projections of future Medicare costs are extremely sensitive to the number and expected needs of potential beneficiaries. While much attention has focused on efforts to reduce Medicare costs by reforming reimbursement procedures and changing the health care delivery system, there are limits to the savings available from these devices. Moreover, their fundamentally one-shot character makes their impact on projected future costs much smaller than demand-related factors such as health status which cumulate year after year.

This paper surveys some relevant evidence, and presents some new calculations, bearing on the effects of mortality improvements. While the available data permit only tentative conclusions, it appears that reductions in morbidity associated with declining mortality have been counterbalanced by high morbidity rates among marginal survivors. As a consequence, the health needs of elderly persons at given ages have not changed very much.

The paper is divided into five sections. The first analyzes recent trends in mortality among the aged. We document the substantial changes which have occurred in both mortality rates and life expectancy, and consider projections of further progress in reducing the mortality of the elderly. We show that a sizable fraction of the elderly population, particularly at extreme ages, is comprised of "marginal survivors."

The second section presents a formal model capable of capturing the various effects of health progress on the elderly population. It allows a role for both population heterogeneity and health progress in reducing morbidity, and provides a basis for understanding both the optimistic and pessimistic views. We use the model to examine the effects of recent mortality progress. Using demographic data for the population cohorts born between 1880 and 1910, we find little support for the "life table rectangularization" hypothesis. Rather, progress in reducing mortality is if anything an increasing function of age.

The third section considers several types of evidence on the age-specific health status of the elderly to distinguish between the optimistic and pessimistic views. Since health is many-faceted, we employ several different measures of health status. The recent evolution of age-specific Medicare costs, institutionalization rates, and disability suggests that the effects of additional survivorship have been about as large as those of reduced morbidity. Continuation of this balance implies that the future aging of the elderly population is likely to be associated with modest increases in per-capita health care costs. Medical costs are higher for the very old than for those in their sixties and seventies, and the increasing proportion of the elderly who will be in their eighties will tend to raise average expenditures per person over age 65. However, there is little evidence to suggest that age-specific costs are likely to increase substantially.

The fourth section examines some micro-econometric evidence bearing on these issues. We analyze data from the Longitudinal Retirement History Survey. Although the data sample is limited by the absence of anyone over the age of 73,

the LRHS can still shed some light on the two views. We show that even controlling for time until death, age has a substantial effect on rates of retirement and medical costs. This confirms our judgement that the observed relationship between age and morbidity, and age and medical costs, is not a spurious one due solely to an underlying link between health status and time till death. Our conclusion considers the policy implications of our findings, and suggests several directions for future research.

I. Mortality Trends among the Aged

I.A. Historical Experience

The last several decades have witnessed dramatic reductions in mortality among the aged. Life expectancy for women at age 65, after rising by 1.8 years between 1950 and 1970, has risen by 1.5 years between 1970 and 1980. For men, the change between 1950 and 1970 was 0.2 years, while the 1970-80 period saw a 1.1 year increase in life expectancy at age 65. These increases in life expectancy were caused by dramatic reductions in mortality rates at all ages. Between 1950 and 1980 the death rate of male 85 year olds fell from 221 per thousand to 178 per thousand, or nearly twenty percent.

Table 1 describes progress in reducing mortality among the aged over the past forty years. We present data on white men and white women, because demographic data for nonwhites at extreme ages are somewhat unreliable.[1] Several patterns emerge from the mortality data. First, mortality gains have been greater for women than for men, despite convergence between the sexes in patterns of employment, lifestyle, and rates of smoking. Second, about equal percentage mortality gains have occurred at all ages. If anything, the gains have been greater among the very old than among those aged 65 to 75. Despite forecasts of ultimate limits on life expectancy and the rectangularization of the life table, there is no evidence of rising mortality rates as more and more people reach extreme ages. The fragmentary available evidence on mortality rates among those over 90 also supports this conclusion.[2] Third, the pace of mortality reductions has accelerated, with especially rapid progess being made during the 1970's.

Table 1

Historical Reductions in Mortality
and Gains in Life Expectancy

Crude Death Rates (per 1000 persons)

WHITE MEN

Year	Age 55-59	Age 65-69	Age 75-79	Age 85+
1950	18.8	40.7	90.1	221.2
1960	17.8	40.5	87.0	217.5
1970	17.7	40.5	86.9	185.5
1980	13.9	33.2	80.7	178.2

WHITE WOMEN

Year	Age 55-59	Age 65-69	Age 75-79	Age 85+
1950	10.2	25.2	69.9	196.8
1960	8.3	21.5	60.8	194.8
1970	8.3	19.2	53.5	159.8
1980	7.2	16.5	45.6	140.4

Life Expectancies

WHITE MEN

Year	Age 55	Age 65	Age 75	Age 85
1950	19.1	12.8	7.8	4.4
1960	19.5	13.0	7.9	4.4
1970	19.5	13.0	8.1	4.6
1980	21.2	14.2	8.8	5.0

WHITE WOMEN

Year	Age 55	Age 65	Age 75	Age 85
1950	22.6	15.0	8.9	4.8
1960	23.8	15.9	9.3	4.7
1970	24.9	16.9	10.2	5.5
1980	26.5	18.5	11.5	6.3

Notes: Death rates are drawn from U.S. Public Health Service, Vital Statistics of the United States, Volume II -- Mortality, Part A, for various years. Life expectancies are from Vital Statistics of the United States, 1980, Life Tables, Volume II, Section 6.

The second part of the table illustrates the substantial changes in life expectancy during this period. For women aged 75, life expectancy increased from 8.9 to 11.5 years between 1950 and 1980, an increase of 29 percent. The gain for men at age 75 was only one year during this period. Even at age 85, there were dramatic improvements: a 14 percent increase for men, and a 31 percent gain for women. In 1950, an 85-year-old woman's life expectancy was only 0.4 years greater than that for an 85-year-old man. In 1980, the difference was 1.5 years.

The causes of these dramatic declines in mortality are not well established. The bulk of the increase in life expectancy at older ages appears to be the result of reductions in mortality rates from cardiovascular diseases. The most recent report of the Social Security Actuary[3] indicates that between 1968 and 1980, mortality from heart disease among persons 65-69 declined at a 2.9 percent annual rate, and mortality from vascular disease fell at a 5.0 percent rate. Partially reflecting these changes, cancer death rates actually rose at a .5 percent annual rate over the same interval.

The cause of these mortality reductions is far from clear. One possible explanation is improvements in access to medical care due to the enactment of Medicare and Medicaid in 1965. The timing of the acceleration in the decline in mortality rates supports this possibility. An alternative possibility is improvements in medical procedures for treating hypertension and heart attacks. Still another possible cause of the decline is improvements in diet and exercise among the aged. In all likelihood the decline in mortality can be traced to some combination of all these factors.

Declines in mortality have potentially important effects on the composition of the aged population. Most obviously, the average age of the elderly population will increase as more and more people survive to older ages. More subtly but probably more importantly, reduced mortality means that the population at any given age will include marginal survivors who will be less healthy and less self sufficient than the remainder of the population.

The magnitude of the mortality declines is well conveyed by the information in Table 2. The table presents estimates of the fraction of 1980 population at various ages who, conditional upon reaching age 50, would not have been alive had they faced the mortality rates of cohorts born 10, 20, 30 or 40 years earlier. We focus on persons who would have reached age 50 to highlight changes for the elderly, and to avoid contaminating our results with changes in infant mortality or other factors affecting younger persons.

The shares of "marginal survivors" are calculated using cohort life tables for persons born in the first year of each decade between 1850 and 1910.[4] These data should be distinguished from those in <u>synthetic</u> life tables, the type commonly used in calculations of life expectancies. In a synthetic table for year t, the death rates at each age correspond to the probability that a person of that age would die in year t. The death rates for each age therefore correspond to different birth cohorts. In a cohort life table, a single birth cohort is followed throughout its life.

We calculated the number of marginal survivors as follows. Let q_t^k denote the probability that a person born in year k dies between birthdays t and t+1, conditional on living to age t. The probability of living to age t in birth cohort k, conditional on reaching age 50, is therefore

Table 2

The Importance of Marginal Survivors: 1980

Fraction of Persons Alive in 1980 Who Would
Not Have Been Alive if Born m Years Earlier

WHITE MEN

Number of Years Born Earlier(m)	Age 60	Age 70	Age 80	Age 90
10	.039	.034	.050	.162
20	.038	.060	.080	.363
30	.054	.085	.151	.584
40	.065	.105	.251	.587

WHITE WOMEN

	Age 60	Age 70	Age 80	Age 90
10	.009	.029	.108	.304
20	.014	.080	.222	.530
30	.036	.144	.351	.723
40	.061	.207	.462	.762

Sources: Authors' calculations based on mortality rates provided by Metropolitan Life Insurance Company, and supplemented with data from the Vital Statistics of the United States, Volume II-Mortality, Part A, for years since 1974.

$$S^k(t) = \prod_{i=50}^{t-1} (1-q_i^k).$$

For persons of age a in 1980, a fraction $S^{1980-a}(a)$ of the members of the birth cohort who reached age 50 are still alive. If these persons had been born m years earlier, the comparable fraction would have been $S^{1980-a-m}(a)$. The proportion of the 1980 population which is accounted for by marginal survivors relative to the cohort m years earlier, MS(a,m), is therefore:

$$MS(a,m) = \frac{[S^{1980-a}(a) - S^{1980-a-m}(a)]}{S^{1980-a}(a)}.$$

The results of these calculations are shown in Table 2. Particularly at old ages, the share of marginal survivors in the population is very high. Fifteen percent of the 80 year old men alive in 1980, and thirty-five percent of the 80 year old women, would have reached age 50 but not have been alive at 80 given the mortality experience of the cohort that preceded them by 30 years. The share of marginal survivors rises rapidly with age. For women at age 60, it is only 0.9 percent, rising to 2.9 percent at age 70, 10.8 percent at 80, and then at a rate of about 4 percentage points per year for each year in the eighties.

The dramatic importance of marginal survivors at extreme ages may be somewhat misleading, since the number of individuals alive at these ages is much smaller than those at earlier ages. We therefore calculated the fraction of the over-50 population which would not have been alive if everyone had faced the life table of the cohort thirty years before them. Over nine percent of the men and 16.9 percent of the women over age 60 in 1980 were marginal survivors.

If people had faced the life table of those born 40 years earlier, twelve percent of the men and twenty-two percent of the women would not have reached their current ages.[5]

Changes in mortality rates or other indicators of health status for the very old are difficult to interpret. The composition of the population has changed quite dramatically through time. Absent general improvements in health, we would expect the large number of marginal survivors to reduce indices of health status at any given age. Of course, the dramatic reductions in mortality could have been accompanied by progress in lowering morbidity rates as well. We consider this possibility in Section III. First, we consider some implications of continuing reductions in mortality.

I.B. Future Trends

Demographic forecasts are notoriously difficult. The substantial mortality gains of the last decade were largely unforecasted. Most observers expected a levelling off in the rate of decline of death rates among the elderly. Nonetheless, it is useful to consider the potential effects on the future population of continued mortality reductions. We rely on the Social Security Administration's Office of the Actuary, which computes three alternative scenarios reflecting different degrees of optimism about future mortality reductions.[6] Alternative II assumes the continuation of current trends, with a gradual adjustment to moderate rates of mortality progress. Alternatives I and III respectively consider slower and faster progress in reducing mortality.

Table 3 displays information on projected death rates. If current

Table 3

Projected Reductions in Mortality, 1980-2080

DEATH RATES PER 1000

Age/Sex Group	Actual 1982	Projected 2000	Projected 2040	Projected 2080
Male 60-64	20.8	16.2	13.8	11.9
Male 65-69	33.7	28.0	23.8	20.6
Male 70-74	49.0	41.4	35.2	30.2
Male 75-79	74.8	63.9	53.9	46.2
Male 80-84	106.2	90.7	76.2	64.9
Male 85-89	158.4	135.1	112.5	95.2
Male 90-94	225.8	191.6	158.6	131.8
Female 60-64	11.2	9.7	8.2	7.0
Female 65-69	16.9	14.8	12.5	10.6
Female 70-74	25.3	20.5	17.1	14.4
Female 75-79	41.1	31.2	25.5	21.1
Female 80-84	65.6	49.0	39.5	32.3
Female 85-89	112.2	84.8	67.6	54.7
Female 90-94	177.6	140.9	111.1	88.9

Source: U.S. Department of Health and Human Services, Social Security Administration, Office of the Actuary, Actuarial Study No. 92, Social Security Area Population Projections, 1984. Projections make Alternative II mortality progress assumptions.

trends continue, the death rate at age 75 for women is expected to decline 24 percent by 2000. At this point, it will be fully 45 percent below its 1960 level. By 2080, the death rates for women at all ages over 75 are projected to be approximately half their current level. For men, progress is less dramatic but still implies a forty percent mortality reduction at high ages.

Table 4 shows the movements in life expectancy at age 65 which these projections imply. We report the forecasts under all three mortality scenarios. Dramatic improvements are clearly a possibility. Under Alternative II, the life expectancy for men at age 65 will rise by 42 percent, to nearly 20 years, by 2040. For women, even the pessimistic projections suggest life expectancies at 65 in excess of twenty years by 2040. The optimistic scenario suggests values of more than 25 years by 2040, and over 29 years by 2080.

These reductions in mortality have important implications for the size and structure of the aged population. In 1982, there were 27.5 million persons in the 65+ age category; this constituted 11.4 percent of the total population, and 19.8 percent as many people as the the population aged 20-64. By 2040, the Alternative II projections imply that the 65+ age group will include 68.8 million persons, or 21.1 percent of the total population and 39 percent as many as the 20-64 age group. The average age of those over age 65 will also rise. If the optimistic projections are accurate, the share of the aged population which is over 85 will rise from 9.7 percent in 1985 to 16.2 percent in 2040. The population at older ages will contain many marginal survivors. Of the 65 year olds alive in 1965, 23.8 percent of the men and 44.6 percent of the women were still alive in 1985. The intermediate projections of

Table 4

Projected Changes in Life Expectancy
at Age 65, 1980-2080

	YEAR			
	1980	2000	2040	2080
Men				
Projection I	14.0	14.8	15.6	16.4
Projection II	14.0	15.7	17.1	18.5
Projection III	14.0	16.6	19.9	23.3
Women				
Projection I	18.4	19.5	20.6	21.5
Projection II	18.4	20.7	22.5	24.3
Projection III	18.4	21.8	25.5	29.1

Source: U.S. Department of Health and Human Services, Social Security Administration, Office of the Actuary, Social Security Area Population Projections, 1984.

the Social Security Actuary suggest that 39.6 percent of those between 65 and 70 in 2000 will be alive in 2020. Less dramatic increases in the proportion of 65 year olds living to be 75 can also be projected.

These data suggest that progress in reducing mortality is having and will have an important impact on the composition of the aged population. These effects are potentially important because there are great differences among the aged in the medical and institutional resources they require. In 1982, the most recent year for which data are available, medical expenditures per capita for persons over 85 were about twice as great as those for persons between 65 and 66.[7] The rate of institutionalization was 11.3 times as great for men over age 85 as for those between 65 and 74.[8] For women, the comparable ratio was 14.6. These figures suggest that the dependency burden of the elderly population could increase substantially with time. It also seems reasonable to expect that the health status of marginal survivors will be worse than that of the remainder of the population.

The adverse effect of increased survivorship on the health status of the elderly population may of course be offset by improvements in our ability to treat chronic illness. The next section presents a formal framework for thinking about the effects of reduced mortality on the health status of the population. In the last two sections, we examine the relative importance of improvements in our ability to manage chronic illness and the changing composition of the population in determining the health status of the elderly.

II. A Formal Model of Mortality Reductions

The interactions between progress in reducing mortality and the health status of the surviving population are complex. On the one hand, measures which lower mortality may also improve health status. Reductions in smoking, improvements in diet, and improved control of hypertension probably improve health at all ages. On the other hand, reductions in mortality may also raise morbidity by changing the composition of the surviving population. An obvious example is provided by those whose lives have been extended through the widespread availability of kidney machines. Mortality reductions also raise morbidity by increasing the average age of the population.

The relative importance of these two effects has been the subject of some dispute. Victor Fuchs[9] and Richard Fries[10] take the optimistic view that health progress is likely to be associated with reduced morbidity. Other authors, notably Manton[11], take the opposite view and suggest that the burden of caring for the elderly population will rise as mortality falls. Which view is correct depends on the source of mortality reductions. Kidney machines and exercise programs will differ in their effects on the health status of the aged population. Ideally, an analysis of recent trends would focus on the differential sources of reduced mortality. However, it is notoriously difficult to isolate the reasons for declining mortality among the elderly. We therefore present a general framework which formalizes the effects of lower mortality rates on health status.[12] This first part of the section presents a theoretical model of mortality in heterogeneous populations, and the second part applies it to analyze the gains in mortality during the last forty years.

II.A. Population Heterogeneity and Mortality

We present a model due to Vaupel, Manton, and Stallard[13] (VMS) which permits decomposition of observed changes in death rates into a component due to health progress which affects all individuals, and a component due to the changing average frailty of the population. The model is stylized in assuming that each individual is endowed with a "frailty" at birth which remains constant throughout life. However, it successfully captures the notion that health progress which reduces the risk of death for all individuals will raise the average frailty of the surviving population, especially at very advanced ages. This composition effect may partly mask mortality improvements.

We assume that the force of mortality for individual i, in cohort j, at age t, $\mu_i^j(t)$, is the product of two terms:

$$\mu_i^j(t) = z_i * \mu^j(t) \tag{1}$$

where $\mu^j(t)$ equals the cohort-specific force of mortality for persons of age t, and z_i is person i's frailty at birth.[14] The force of mortality and the age-specific death rate, $q_i^j(t)$, are linked by the approximation

$$\begin{aligned} q_i^j(t) &= \Pr(\text{person } i \text{ in cohort } j \text{ dies between ages } t \text{ and } t+1, \\ &\qquad \text{conditional on reaching age } t) \\ &= -\log(1 - \mu_i^j(t)). \end{aligned} \tag{2}$$

The probability that a type-z_i individual will survive to age m, $S_i^j(m)$, is

$$S_i^j(m) = \exp(-\int_0^m \mu_i^j(t)dt) = \exp(-z_i \int_0^m \mu^j(t)dt). \tag{3}$$

The population force of mortality at each age, $\bar{\mu}^j(t)$, is just a weighted

average of individuals' $\mu_i^j(t)$'s. It depends upon both the distribution of frailties among those who are alive, and the cohort-specific force of mortality $\mu^j(t)$. From equation (1),

$$\bar{\mu}^j(t) = \bar{z}^j(t) \cdot \mu^j(t) \tag{4}$$

where $\bar{z}^j(t)$ equals the mean frailty of survivors in cohort j at age t. The rate of morbidity, $\bar{\nu}^j(t)$, can also be modelled as a function of average frailty and a cohort-specific morbidity function, $\nu^j(t)$:

$$\bar{\nu}^j(t) = \bar{\phi}(\underset{\sim}{z}^j(t)) \cdot \nu^j(t). \tag{5}$$

The ϕ function translates mortality-relevant frailties into morbidity-relevant ones. The cohort-specific morbidity function is designed to capture various factors such as medical progress which affect morbidity.

To make this model operational, we must make some assumption about the distribution of frailties at different ages. We define $f^j(z,t)$ as the probability density function for frailties of individuals in cohort j at age t. Vaupel, Manton, and Stallard assume that frailties at birth follow a gamma distribution. The gamma is sufficiently flexible to allow for a wide variety of distribution patterns. It also has the appealing property that if frailties at birth are gamma distributed, then so are frailties of the survivors at all subsequent ages. We postulate that

$$f^j(z,0) = \lambda_j^k z^{k-1} e^{-\lambda_j z}/\Gamma(k) \tag{6}$$

which is the gamma density with parameters λ_j and k. It's mean is k/λ_j, which

equals one, the average frailty of individuals at birth. This implies $k = \lambda_j$. The variance equals $\sigma_j^2(0) = k/\lambda_j^2$. VMS show the density of frailties for age-t survivors is:

$$f^j(z,t) = [\lambda_j(t)]^k z^{k-1} e^{-\lambda_j(t)z} / \Gamma(k). \tag{7}$$

The parameters of this gamma distribution are $\lambda_j(t) = \lambda_j - \log \bar{S}^j(t)$ and k. $\bar{S}^j(t)$ is the fraction of cohort j surviving to age t. Using (7), the average frailty of age-t survivors is therefore

$$\bar{z}^j(t) = k/[\lambda_j - \log \bar{S}^j(t)] \tag{8}$$

and the variance in frailties at age t is

$$\sigma_j^2(t) = k/[\lambda_j - \log \bar{S}^j(t)]^2. \tag{9}$$

Using the fact that $\lambda_j = k$, the mean frailty at age t may be rewritten as

$$\bar{z}^j(t) = k/[k - \log \bar{S}^j(t)]. \tag{10}$$

Average frailty declines as a cohort ages, since death is more likely to remove frailer members of the population at earlier ages. The variance of frailties is also a declining function of age. This is intuitively reasonable, since at very advanced ages only the strongest members of the original population, those with the lowest z's, will remain alive.

The importance of accounting for heterogeneity rises with age. This can be illustrated by considering a reduction in mortality which lowers the cohort specific force of mortality by a constant fraction (δ) at all ages:

$$\tilde{\mu}(t) = \delta \cdot \mu(t) \tag{11}$$

This mortality improvement will affect the mean frailty of the very old by more than that for other groups, although the measured death rates for this group will be the least influenced by this improvement. This is because the change in mortality at each age has two components:

$$\frac{d\bar{\mu}(t)}{d\delta} = \frac{d(\delta\mu(t)\bar{z}(t))}{d\delta} = \bar{z}(t)\mu(t) + \mu(t)\frac{d\bar{z}(t)}{d\delta}\,\delta\ . \tag{12}$$

The first term yields a reduction in mortality rates as δ falls. It corresponds to the direct reduction in the mortality rate for persons who survive to each age. The second term has the opposite effect; as δ falls, it shows that average frailty at each age will rise, causing some <u>increase</u> in the observed age-specific death rate. At all ages average frailty rises as δ falls, i.e., $d\bar{z}(t)/d\delta$ is less than zero, signalling a mortality improvement. This effect is largest at old ages. Since the direct reduction in mortality rates is a constant proportion at all ages, the <u>observed</u> response to an improvement such as (11) will be smallest at high ages. This is what one would expect intuitively. The selection effect of mortality improvements cumulates through time, and so has its greatest impact at high ages.

A central insight which follows from models of mortality in heterogeneous populations is that the mortality prospects facing an individual may be poorly described by observed population death rates. Because selection operates to increase the fraction of stronger individuals in the population as it ages, the life expectancy which one would estimate from the actual survival curve will always overstate the life expectancy for an average individual at birth. The

relationship between the force of mortality for a type-z_i person at age t, and the observed cohort force of mortality at that age, is

$$\mu^j(z_i,t) = \bar{\mu}^j(t)*z_i*(\bar{S}^j(t))^{-1/k} \quad (13)$$

where $\bar{S}^j(t)$ is the fraction of cohort j which has survived to age t.[15] This shows that the divergence between the risks facing an individual of any given frailty, and those observed to face the population, rises with age since $\bar{S}^j(t)$ declines. It follows that observed progress in reducing mortality will tend to understate true progress in reducing mortality at old ages. This is because the aged population following a reduction in mortality contains marginal survivors whose experience is likely to be less favorable than that of other members of the population.

Models of population heterogeneity similar to the one described above have often been proposed as potential explanations of the "crossover effect," the observation for example that mortality rates for very old black women in the United States are lower than those for whites, although at earlier ages the pattern is just the opposite. Crossovers are possible when populations are comprised of heterogeneous individuals, since high mortality rates in early life may select a very hardy pool of survivors. A growing body of evidence, however, suggests that observed crossovers may be spurious artifacts of poor demographic data and inaccurate age reporting. Coale and Kisker[16] demonstrate a strong correlation between the frequency of "age heaping," the tendency for ages which end in five or zero to attract a higher fraction of census responses than the adjoining ages, and the presence of cross-over effects in a sample of demographic data for 20 countries. Even if heterogeneity is not the full expla-

nation for crossovers, however, models such as the one described here could capture an important aspect of health progress.[17]

The model which we have sketched imposes several strong restrictions on the nature of mortality reductions. For example, a reduction in the cohort's baseline mortality, $\mu^j(t)$, affects very frail individuals much more than those who were initially healthy. There are undoubtedly some forms of medical progress which affect healthier individuals more than those who are extremely frail, and it would be desirable to allow for such progress in a more general framework. There also may be errors introduced by our assumptions about the functional form of the frailty distribution, although it is difficult to assess their impact. Finally, when we make comparisons across cohorts born at different ages, we assume that the average frailty at birth is the same for each cohort. This seems a natural starting point, although further work might examine the extent to which changing patterns of neonatal care could influence the value of $\bar{z}^j(0)$ across cohorts.

II.B. Application to Recent Mortality Gains

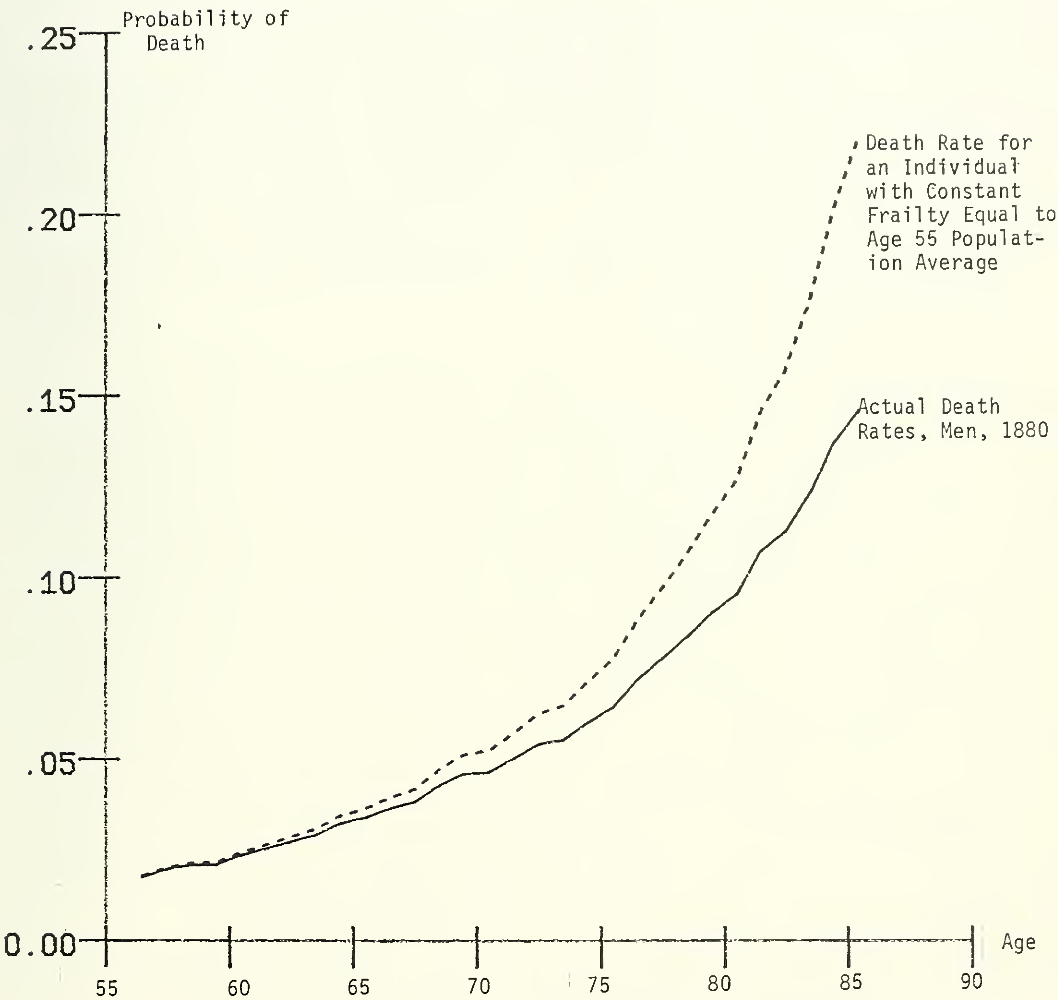
When we normalize the mean cohort frailty to unity at birth, parameterizing the frailty distribution then reduces to the problem of choosing k. The variance of frailties at birth is equal to 1/k. As k rises, the dispersion of frailties declines until in the limiting case of $k=\infty$, there are no differences among cohort members.

Manton, Stallard, and Vaupel[18] analyze mortality data on the cohorts of white men and white women born in the United States in five year intervals between 1850 and 1880. They estimate that k equals 3.93 for the male population, and 2.84 for women. They obtain similar results using Swedish data. We

attempted our own estimates of the variance in frailty among the aged population by trying to fit the observed cohort life tables, assuming a mixture of individual Gompertz mortality functions. These efforts failed to yield reasonable estimates. The results suggested a complete absence of heterogeneity. This may have resulted from our assumption about the functional form of the individual hazard functions. It would be valuable to reconsider the estimation in future research.

Our analysis assumes k=4.0 for both men and women; this probably overstates k and understates the dispersion of frailty. The parameters imply that 81.5 percent of men at birth have frailties between one-half and twice the average frailty. To illustrate the difference between the population mortality experience and that facing an individual of constant frailty, we compute the probabilities of dying each year for a man who was born in 1880 with a frailty of 0.56, the average frailty for those who survived to age 55. We then compare the probabilities that he will die in each year after age 55 with the observed cohort death rates for these ages. The two sets of death rates, denoted $\bar{q}^j(t)$ and $q^j(.56,t)$ are plotted on Figure 1. As time elapses after age 55, the difference between the constant frailty individual's probability of dying and that for the cohort as a whole widens. At age 60, for example, the observed death rate is 0.237, while that for our constant-z individual is 0.245. By age 75, the difference is more dramatic: 0.645 versus 0.778. These trends reflect the declining average frailty of the surviving population; our constant-frailty person is increasingly among the frailest members of the surviving cohort. At age 60, he is frailer than 58.3 percent of the suriving cohort; by age 75, 68.3 percent are less frail.

Figure 1: Death Rates for Cohort and Constant-Frailty Person



We use the Vaupel-Manton-Stallard technique to analyze changes in mortality rates and life expectancy for individuals born in 1880 and 1910. First, we calculate the observed change in mortality rates and compare it with the change which would have taken place assuming that the average frailty of survivors in the 1880 cohort had applied to similar-aged survivors in the 1910 cohort. We also compute the changes in life expectancy at each age between the two cohorts, again making corrections for movements in average frailty. This enables us to identify the ages at which substantial mortality gains have taken place.

Computing life expectancies requires data on the probability of death at ages up to 100 because the exact age at which people die is important. Unfortunately, the maximum age reported in our cohort life tables is 85. For the 1910 cohort, data are available on persons up to 72 years of age. We extended our tables to age 100 by fitting a Gompertz curve, a standard functional form relating age and the force of mortality,[19] to our data on each cohort's death rates at ages between 55 and 85. We then use this curve to predict values of death rates at ages greater than 85. The Gompertz curve specifies that the force of mortality rises exponentially over time: $\mu^j(t) = \mu^j(a)e^{\beta(t-a)}$ where a in our estimates equals 55. This specification implies simple regression models for mortality rates:

$$\log\left(-\log[1-q^j(t)]\right) = \alpha_j + \beta_j*(t-50) + \epsilon_{jt} . \qquad (15)$$

The results of our estimates for each cohort are shown in the appendix.

The probabilities of death which would have been observed for the 1910

cohort had average frailty at each age equalled that for the 1880 are:

$$\tilde{q}^{1910}(\bar{z}^{1880}(t),t) = 1 - \exp[-k*\log(1-\bar{q}^{1910}(t))*\bar{z}^{1880}(t)/\bar{z}^{1910}(t)]. \quad (14)$$

The expression depends upon $\bar{z}^{1910}(t)$, the average frailty at age t for the survivors in the 1910 cohort, and $\bar{z}^{1880}(t)$, the average frailty for survivors from the 1880 cohort when they were t years old. The results of these calculations are shown in Table 5, which reports the actual $\bar{q}^{1880}$ and $\bar{q}^{1910}$ and the frailty-adjusted death probabilities at five year intervals for both men and women. The table shows that for men, the probability of dying at age 65 declined from 0.0341 to 0.0291 between the 1880 and 1910 cohort, a decline of 0.0050. At age 80, the decline was more pronounced, from 0.0954 to 0.0578. The table also shows that the decline would have been even larger at both ages if the average frailty of the respective populations had remained constant at their 1880 level. The change at age 65 would have been 0.0067, while at age 80, the constant-frailty decline in death rates equals .0430. Figure 2 shows the reductions in death rates, with and without our frailty adjustment, for all ages between 55 and 85. The figure vividly demonstrates that the largest reductions in mortality occured at very advanced ages. While a male survivor's probability of dying falls by nearly one quarter at age 55, it is reduced by roughly fifty percent at all ages above 80.

The table shows that even more pronounced changes have occurred for women. At age 65, the observed mortality rates declined from .0229 to .0135 between 1945 and 1975, the dates when women in the 1880 and 1910 cohorts turned 65. Adjusting for changes in frailty yields a relatively small additional improvement, converting the 1975 mortality rate to .0125. At older ages, the

Table 5
Changes in Actual and Frailty-Adjusted Mortality Rates, 1880-1910 Birth Cohorts

Actual Probabilities of Death, Men

Age	$\bar{q}^{1880}$	$\bar{q}^{1880}-\bar{q}^{1890}$	$\bar{q}^{1890}-\bar{q}^{1900}$	$\bar{q}^{1900}-\bar{q}^{1910}$	$\bar{q}^{1910}$
55	1.65	0.09	0.21	-0.01	1.36
60	2.37	0.16	0.14	-0.03	2.10
65	3.41	0.25	0.02	0.23	2.91
70	4.66	0.13	-0.11	0.84	3.80
75	6.45	-0.11	0.12	1.54*	4.90*
80	9.54	0.04	0.61	3.11*	5.78*
85	14.61	1.80	2.00*	4.15*	6.66*

Probabilities of Death Assuming 1880 Frailty Levels, Men

Age	$\bar{q}^{1880}$	$\bar{q}^{1880}-\tilde{q}^{1890}$	$\tilde{q}^{1890}-\tilde{q}^{1900}$	$\tilde{q}^{1900}-\tilde{q}^{1910}$	$\tilde{q}^{1910}$
55	1.65	0.11	0.25	0.02	1.28
60	2.37	0.19	0.20	0.01	1.98
65	3.41	0.30	0.11	0.27	2.73
70	4.66	0.20	0.02	0.89	3.55
75	6.45	-0.02	0.29	1.64*	4.54*
80	9.54	0.17	0.85	3.28*	5.24*
85	14.61	2.02	2.40 *	4.43*	5.76*

Actual Probabilities of Death, Women

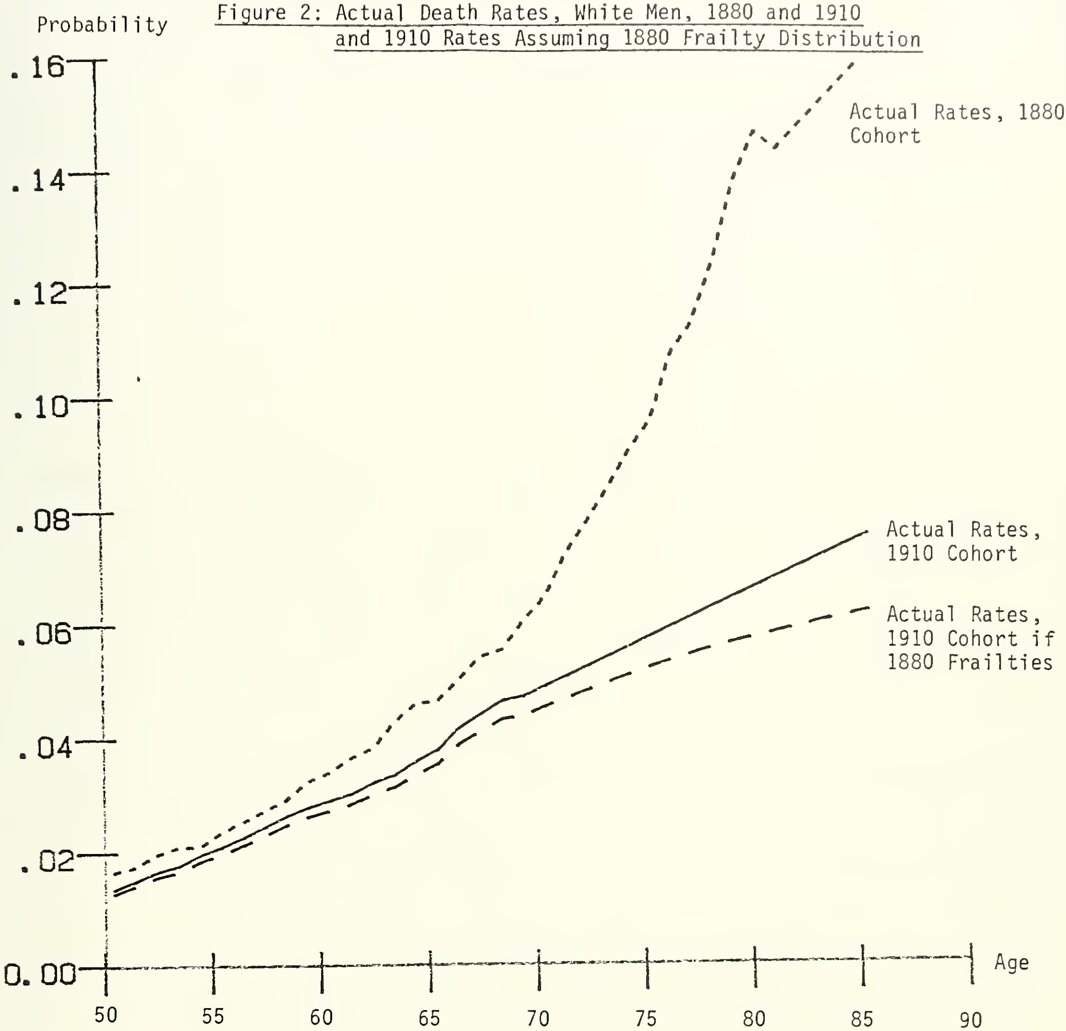
Age	$\bar{q}^{1880}$	$\bar{q}^{1880}-\bar{q}^{1890}$	$\bar{q}^{1890}-\bar{q}^{1900}$	$\bar{q}^{1900}-\bar{q}^{1910}$	$\bar{q}^{1910}$
55	1.17	0.23	0.25	0.04	0.65
60	1.57	0.35	0.23	0.04	0.95
65	2.29	0.54	0.23	0.17	1.35
70	3.07	0.45	0.25	0.51	1.86
75	4.50	0.48	0.44	0.93*	2.65*
80	7.21	0.94	0.94	2.19*	3.14*
85	11.98	2.45	2.94*	2.96*	3.63*

Probabilities of Death Assuming 1880 Frailty Levels, Women

Age	$\bar{q}^{1880}$	$\bar{q}^{1880}-\tilde{q}^{1890}$	$\tilde{q}^{1890}-\tilde{q}^{1900}$	$\tilde{q}^{1900}-\tilde{q}^{1910}$	$\tilde{q}^{1910}$
55	1.17	0.25	0.27	0.05	0.61
60	1.57	0.37	0.26	0.06	0.88
65	2.29	0.58	0.28	0.19	1.25
70	3.07	0.52	0.33	0.52	1.70
75	4.50	0.61	0.56	0.94*	2.39*
80	7.21	1.17	1.14	2.15*	2.76*
85	11.98	2.90	3.18*	2.88*	3.02*

Source: Authors calculations based on unpublished cohort life table data for while men and while women provided by the Metropolitan Life Insurance Company. The $\bar{q}$ series are actual cohort death rates, while $\tilde{q}$ are the death rates which would have been observed if the average frailty of the 1880 cohort had prevailed at all times. Starred values are based or extrapolated death probabilities. See text for details of calculations.

Figure 2: Actual Death Rates, White Men, 1880 and 1910 and 1910 Rates Assuming 1880 Frailty Distribution



frailty adjustment matters somewhat more. At age 85, for example, observed death probabilities fell from .1198 to .0363, while keeping the 1880 cohort's frailty level, $\bar{q}^{1910}$ would have fallen to .0302.

Studying changes in death rates is one way to identify the ages at which the most progress has been made against mortality. However, the claim that substantial gains have occured at extreme ages may be of little significance if the fraction of the population which lives to these ages is trivial. An alternative measure of where gains have been made is the change in life expectancy at a given age between two cohorts.

Our life expectancy calculations are shown in Table 6. The table shows the actual life expectancies at ages 55 through 85 for individuals born in 1880, 1890, 1900, and 1910. The changes in life expectancy for both men and women occur disproportionately between the last two cohorts. For example, a man born in 1880 who lives to age 75 has a life expectancy of 8.3 years. If he had been born in 1900 and reached age 75, his life expectancy would be 9.3 years. However, an individual born just ten years later, in 1910, has a life expectancy of 12.45 years at age 75. The absolute gain in life expectancy is a smoothly declining function of age, while the percentage gain rises with age.

The difference between actual and frailty-adjusted life expectancies can be seen by comparing the first and second panels of the table. If the average frailty of the 1910 cohort at each age had equalled the same-age average frailty of the 1880 cohort, life expectancy would have been roughly one year greater. At age 55, for example, it would have raised male life expectancy from 22.4 to 23.5 years. At age 80, changes from 10.66 to 11.60 years from men, and

Table 6

Changes in Actual and Frailty-Adjusted Life Expectancies, 1880-1910 Birth Cohorts

Actual Life Expectancies, Men

Age	$\bar{e}^{1880}$	$\bar{e}^{1890}-\bar{e}^{1880}$	$\bar{e}^{1900}-\bar{e}^{1890}$	$\bar{e}^{1910}-\bar{e}^{1900}$	$\bar{e}^{1910}$
55	19.06	0.46	0.59	2.29	22.41
60	15.87	0.36	0.51	2.55	19.29
65	13.03	0.25	0.50	2.87	16.65
70	10.56	0.25	0.53	2.98	14.32
75	8.31	0.33	0.70	3.10*	12.45*
80	6.52	0.45	0.84	2.84*	10.66*
85	5.44	0.57	0.61*	2.21*	8.83*

Life Expectancies Assuming 1880 Frailty Levels, Men

Age	$\bar{e}^{1880}$	$\tilde{e}^{1890}-\bar{e}^{1880}$	$\tilde{e}^{1900}-\tilde{e}^{1890}$	$\tilde{e}^{1910}-\tilde{e}^{1900}$	$\tilde{e}^{1910}$
55	19.06	0.61	0.93	2.85	23.45
60	15.87	0.50	0.82	3.12	20.31
65	13.03	0.38	0.80	3.44	17.66
70	10.56	0.37	0.81	3.57	15.31
75	8.31	0.45	0.96	3.71*	13.42*
80	6.52	0.57	1.10	3.42*	11.60*
85	5.44	0.70	0.84*	2.68*	9.66*

Actual Life Expectancies, Women

Age	$\bar{e}^{1880}$	$\bar{e}^{1890}-\bar{e}^{1880}$	$\bar{e}^{1900}-\bar{e}^{1890}$	$\bar{e}^{1910}-\bar{e}^{1900}$	$\bar{e}^{1910}$
55	22.55	1.94	2.14	3.48	30.12
60	19.01	1.71	1.97	3.60	26.28
65	15.75	1.43	1.85	3.72	22.75
70	12.76	1.27	1.74	3.68	19.45
75	10.04	1.16	1.77	3.68*	16.65*
80	7.78	1.15	1.77	3.21*	13.92*
85	6.37	1.01	1.43*	2.30*	11.11*

Life Expectancies Assuming 1880 Frailty Levels, Women

Age	$\bar{e}^{1880}$	$\tilde{e}^{1890}-\bar{e}^{1880}$	$\tilde{e}^{1900}-\tilde{e}^{1890}$	$\tilde{e}^{1910}-\tilde{e}^{1900}$	$\tilde{e}^{1910}$
55	22.55	2.30	2.66	3.89	31.40
60	19.01	2.06	2.46	4.00	27.54
65	15.75	1.78	2.34	4.10	23.96
70	12.76	1.61	2.21	4.03	20.61
75	10.04	1.49	2.21	3.98*	17.72*
80	7.78	1.46	2.18	3.45*	14.87*
85	6.37	1.30	1.75*	2.43*	11.86*

Source: Authors calculations based on unpublished cohort life table data for white men and white women provided by the Metropolitan Life Insurance Company. The $\bar{e}$ series are the actual life expectancies for each cohort, while the $\tilde{e}$ are those which would have prevailed if average frailty at each age had remained constant at the 1880 cohort levels. Starred values are based on death rate extrapolations. See text for details of calculations.

13.92 to 14.87 years for women, would be observed. Put another way, the marginal survivors at each age have lower life expectancies than those who would have lived to that age in the previous cohort. Since the difference between the actual and frailty-corrected estimates is approximately the same at all ages, the proportionate change induced by the frailty correction is largest at old ages.

These data cast doubt on the view of Fries and Crapo[20] and others that longevity is currently pushing up against an upper limit. Taking account of heterogeneity, it appears that life expectancy is increasing more at old than young ages. It is increasing more for women than for men, even though women already have longer life expectancies. We find little evidence to confirm the view that mortality is increasingly bunched at some specific age. Apparent evidence of life curve rectangularization could be a manifestation of heterogeneity.

The results in this section suggest that if the framework sketched here is roughly accurate, the changing frailty mix between the 1880 and 1910 cohort should have reduced the gains in both mortality rates and life expectancies which would have taken place under if the survivors were of the same frailty as those in the 1880 cohort. These substantial changes in average frailty should also have had other effects. If a larger fraction of the population survives until age 65, then the variance of frailties at that age will increase. This should imply an increase in the variance of longevities after age 65. For men reaching age 65 in 1925, the variance of the remaining years of life was 40.7; for those reaching 65 in 1965, it was 66.5. The comparable

figures for women in these two years were 45.2 and 87.0, respectively. This accords with the predictions of the heterogeneity model.

A second prediction of this model is that falling mortality rates should have lowered the average health status of the population at advanced ages. Health status is difficult to describe using a single quantitative measure. In the next section we see if any dramatic changes are suggested by several kinds of evidence of health status.

III. Evidence on Changing Morbidity and Health Needs

While it is difficult to assess the extent of heterogeneity in the population, the analysis in the preceding section suggests that changes in mortality rates could result in important changes in the composition of the surviving population. If this were the only force acting on the health status of the aged, one would expect to see substantial deterioration, especially among those at high ages. This supports the pessimistic view of health progress. Another explanation could also be proposed to account for reductions in health status. Episodes of morbidity may weaken individuals. Recent progress may have raised the threshold below which an individual's resilience may fall without causing death. If illness-induced reductions in resilience are persistent, then health progress may raise the average frailty (lower resilience) of the surviving population.

Victor Fuchs' recent optimistic analysis of changes in health status[21] does not consider the changing composition of the aged population. Rather, he focuses on the effects of broadly-defined health progress on the health status of a given aged person. He proposes an intruiging model for thinking about the linkages between aging and health. He argues that for medical care costs, disability, and institutionalization, age is better measured backwards from death rather than forwards from birth. Medical care costs, for example, have been shown to be highly concentrated in the year or two immediately preceding death. Since death rates at all ages have declined, Fuchs' view would lead to the expectation that the health status of the elderly population should actually be

improving since the average number of years till death is rising. The fraction of the population at each age who are within one or two years of death has declined, as evidenced by changing death rates. Fuchs' view predicts that one should see greater improvements in health status at old ages than at younger ages because of the greater absolute reduction in death rates among the elderly. Any other view emphasizing the importance of medical developments or changes in lifestyle in reducing morbidity would also lead one to expect trend improvements in the health of the elderly.

These two views thus offer dramatically opposite predictions about trends in the health status of the elderly population and about the relative health status of persons at different ages. One cannot doubt the existence of both positive developments which reduce morbidity, and changes in the composition of the population which tend to increase illness and disability. The central question is which effects predominate. To investigate this issue, we examine a number of indices of the health status of the elderly. When possible, we look at age-specific measures to avoid biases due to the aging of the population.

Table 7 reports the fraction of the aged population residing in nursing homes for various years between 1963 and 1982. Annual data on the nursing home population are not collected; we report the findings of the five most recent surveys. For both men and women under 85, the data show a steady upwards trend in the rate of institutionalization from 1963 until 1977 and then a small decline in 1982. There is a substantial increase in institutionalization rates at all ages between 1963 and 1969. At least the first part of this increase may reflect increases in the level of public support for nursing home care,

Table 7

Percentage of Population Resident in
Nursing Homes: 1963 - 1977

Age/Sex Group	1963	1969	1973	1977	1982
Men					
65-74	0.68	0.99	1.13	1.27	1.23
75-84	2.91	3.60	4.08	4.74	4.17
85+	10.56	13.08	18.04	14.00	13.93
Women					
65-74	0.88	1.29	1.31	1.59	1.55
75-84	4.75	6.23	7.11	8.06	6.89
85+	17.51	24.76	29.06	25.15	22.69

Source: Data for 1963-1977 are drawn from U.S. Department of Health and Human Services, Public Health Service, Vital and Health Statistics, Series 13 Number 51: Characteristics of Nursing Home Residents: Health Status, and Care Received, 1981, p.4. The 1982 calculations are based on unpublished data from the National Master Facility Survey provided by the National Center for Health Statistics.

especially through Medicaid.[22] The pattern among very old men and women, those over 85, is quite different, however. The rate of institutionalization rises steadily until 1973 and then declines sharply through 1982. This pattern is rather surprising. Non-health determinants of the rate of institutionalization, such as the availability of public assistance or increased viability of remaining at home, would be expected to exhibit similar trends for all age groups. Thus, these data might suggest that the health status of the very old is improving relative to that of their younger counterparts.

An interesting feature of the data is that at all ages women have considerably higher rates of institutionalization than do men despite their longer life expectancy. This probably reflects their much greater likelihood of being widowed. The data also reject popular stereotypes about the pervasiveness of institutionalization. Even for the extreme aged, less than one-fifth of the population is in a nursing home. Only about five percent of the elderly population is institutionalized at any point in time.[23] Unfortunately, it is difficult to draw stong conclusions about the health status of the elderly from institutionalization rates, since much of the variation in this rate may be due to non-health factors.

Information on the pattern of Medicare expenditures on persons of different ages is presented in Table 8. Because the overall level of Medicare spending is driven by economic and non-economic forces beyond the scope of this paper, we focus only on relative levels of expenditure on persons of different ages. The table reports the average Medicare reimbursement per enrollee in each age group, scaled by the average reimbursement per enrollee in the 65-66 age

Table 8

Age-Specific Medicare Expenditure Patterns

Age/Sex Group	1966	1971	1977	1982
White Men				
65-66	1.000	1.000	1.000	1.000
67-68	1.105	1.362	1.049	1.103
69-70	1.152	1.428	1.131	1.189
71-72	1.276	1.518	1.216	1.302
73-74	1.404	1.593	1.338	1.447
75-79	1.572	1.763	1.477	1.615
80-84	1.742	2.009	1.640	1.846
85+	1.972	2.086	1.865	1.958
White Women				
65-66	1.000	1.000	1.000	1.000
67-68	1.199	1.342	1.029	1.098
69-70	1.180	1.432	1.100	1.212
71-72	1.301	1.567	1.176	1.292
73-74	1.398	1.725	1.324	1.412
75-79	1.584	1.912	1.522	1.640
80-84	1.804	2.207	1.774	1.927
85+	2.019	2.243	1.974	2.096

Notes: Each column shows the ratio of Average Medicare Expenditures per Enrollee to expenditure per enrollee in the 65-66 age category. Data are based on authors' calculations, using information drawn from annual issues of Medicare Program Statistics, and Health Care Financing: Program Statistics published by the Health Care Financing Administration.

group. The data show very little variation over time in the age pattern of Medicare expenditures. The ratio of expenditures on enrollees over 85 to those between 65 and 66 was 1.97 for white men in 1966, and 2.02 for white women. In 1982, the comparable values were 1.96 and 2.10, respectively. Similar patterns emerge in other years and at intervening ages, with one exception. To some extent, 1971 is an outlying year. This appears to be entirely due to unusually low expenditures on persons aged 65 to 66. Relative levels of expenditures between other ages are not out of line with the data for other years. Movements in the fraction of enrollees who receive some medical services also show a similar pattern. In 1966, 30.2 percent of Medicare enrollees between 65 and 66 years of age received service, compared with 48.2 percent of those over age 85. In 1982, the rates were 57.5 percent and 73.3 percent, respectively. The growth in utilization was larger for those at younger ages than for the very old.[24]

Data on both utilization and care levels suggest that the effects of health progress and the changing composition of the population have offset each other. They do not support the view that health progress inevitably carries with it huge expenditure burdens for marginal survivors, as the heterogeneity model of the last section would suggest. While close to half of the 1982 population of persons over 85 were marginal survivors, compared to a lower fraction in earlier years, relative Medicare costs remained roughly constant.

Data on Medicare expenditures and institutionalization have the virtue of being objective, but the problem of being influenced by a variety of factors other than the health status of the elderly. Another source of health status information is the extent of activity limitation and disability in the elderly

population. Data on these health measures for the civilian noninstitutional population are collected in the Health Interview Survey, and tabulations are periodically published by the National Center for Health Statistics. Table 9 presents information on the number of restricted activity days per year for subgroups of the elderly population for various years. Because the numbers refer only to the non-institutionalized population, they may be heavily influenced by fluctuations in the institutionalization rate, especially since persons in institutions are likely to have substantial amounts of restricted activity.[25]

In general, the data on restricted activity display no clear trend. The reduced morbidity and frailty-composition effects of medical progress again seem to offset each other. The incidence of restricted activity days for both men and women aged 75-plus is nearly the same in 1980 as in 1961. For the 65-74 age group, there is slight evidence of an increase in disability days when we focus on the 1980 data. This trend is not present in the 1975 survey, however. These data confirm the inference drawn from the data on institutionalization; the health of the elderly has not worsened and may have improved in recent years. In particular, they offer no support for substantial reductions in health at extreme ages where the incidence of marginal survivors is greatest.

Similar conclusions are suggested by information on the incidence of Bed Disability Days in Table 10, which also appear to be relatively constant between 1961 and 1980. For all age groups except women aged 64-74, the incidence of bed disability days was lower in 1980 than in 1961. In some cases, such as women aged 75-plus, the improvement suggests an average reduction of

Table 9

Average Number of Restricted Activity Days per Person, 1961-1980

Year	Men 65-74	Men 75+	Women 65-74	Women 75+
1961	31.9	36.1	34.8	46.2
1963	31.3	41.4	36.3	49.6
1965	30.9	36.0	30.7	41.9
1968	31.2	35.0	30.3	47.6
1971	26.6	38.8	30.6	44.9
1975	31.1	40.7	36.2	49.2
1980	34.2	36.0	39.2	46.6

Source: U.S. Public Health Service, Vital and Health Statistics, Series 10, various issues entitled Disability Days.

Table 10

Incidence of Bed Disability Days Among Person Over Age 65

	AGE/SEX GROUP			
Year	Men 65-74	Men 75+	Women 65-74	Women 75+
1961	11.4	14.6	12.5	23.6
1963	10.8	17.6	11.3	20.0
1965	11.5	14.5	11.1	16.0
1968	12.0	16.2	11.6	20.9
1971	9.0	17.9	10.9	18.9
1975	9.8	17.0	10.6	17.7
1980	10.9	13.4	12.9	19.1

Source: U.S. Public Health Service, Vital and Health Statistics, Series 10, period volumes entitled Disability Days.

nearly one week per year of disability. For men over 75, a substantial improvement in health status is observed between 1975 and 1980, parallelling the observations made above. The principal gains for extremely old women occurred during the 1960s. An interesting aspect of the data is that the disability rate is higher for women than for men, particularly in the over-75 category. This tends to contradict Fuchs' view that time until death is a good indicator of health status, since women have longer life expectancies than men. It is consistent with the heterogeneity view since a much larger fraction of women than men reach the age of 75.

IV. Microeconometric Evidence

The data presented in the preceding section suggest that the health of the elderly population has not changed much through time despite the changing composition of the elderly population and progress in reducing morbidity. This suggests that these two effects have been roughly offsetting. In this section we present some very crude microeconometric evidence bearing on this question. We test Fuchs' hypothesis that "relevant age" should be measured backwards from death rather than forwards from birth by making use of longitudinal data from the Retirement History Survey. The survey includes information on individuals' ages as well as dates of death for the quarter of the original sample which died between 1969 and 1979. We compare the power of the age and time-till-death variables in explaining hospital care costs. Because hospitalization costs are only one indicator of health status, we also compare the performance of the two age variables as predictors of retirement decisions.

The Retirement History Survey suffers from several drawbacks as a source for an investigation of this type. First, it only provides data on the youngest part of the aged population. Even at the end of the sample, the oldest person in the sample was only 73. Second, the information on medical costs in many years refers only to out of pocket expenses. Since individuals become eligible for Medicare at age 65, it is not possible to satisfactorily estimate the effects of chronological aging on health care costs except for those under 65. The existence of other third-party payments also makes these data difficult to interpret. In 1969, however, before any survey participants were eligible for

Medicare, the survey asked for detailed data about hospital costs borne by both the respondent and any third parties. We use only this information in our study of health care expenditures.

We estimate hospital cost equations of the form:

$$HCOST_i = \alpha_0 + \alpha_1*MAR_i + \alpha_2*EDUC_i + \alpha_3*SMSA_i + \alpha 4*RACE_i$$
$$+ \alpha_5*AGE_i + \alpha_6*YTD_i + \alpha_7*SURVIVOR_i + \epsilon_i \ .$$

where MAR is a dummy variable equal to one for persons who were married, EDUC represents years of schooling, SMSA is a dummy variable equal to 1 for persons residing within an SMSA, RACE equals 1 for non-whites, AGE is chronological age, and YTD equals years to death. The survey records both month and year of death, so we are able to measure this variable quite accurately. We also include SURVIVOR, a dummy variable equal to one if the respondent survived until the end of the sample period. We assigned these individuals the maximum value for YTD as well. The dependent variable is total hospital costs during 1968. No economic variables like wealth were included in the equation because of the possibility that they were affected by individual choices based on knowledge of health status.

The results of several specifications of the hospital cost equation are shown in Table 11.[26] Increases in age, and reductions in the number of years to death, both raise hospital costs. A one year reduction in years to death has a slightly larger effect on costs than one additional calendar year of age. The most important finding, however, is that the effects of age and time-till-death are largely independent. Although equations including the time-till-death

Table 11

Reduced Form Hospital Cost Equations, 1969 LRHS Sample

	Equation 1	Equation 2	Equation 3
Constant	-1183.7 (604.6)	193.7 (95.7)	-874.0 (609.7)
MARRIED	70.9 (54.3)	78.7 (54.1)	74.9 (54.1)
EDUCATION	4.3 (8.1)	7.3 (8.1)	6.4 (8.1)
SMSA	57.7 (34.1)	70.5 (34.1)	69.4 (34.1)
RACE	-35.4 (74.3)	-21.4 (74.1)	-25.1 (74.1)
AGE	19.7 (10.3)	---	17.8 (10.1)
YEARS TO DEATH	---	-23.5 (12.1)	-24.5 (12.1)
SURVIVOR	---	-7.7 (68.3)	4.3 (68.5)
SSR	177.7	176.1	175.4
R^2	.012	.021	.025

Notes: All equations are estimated with 786 observations from the Brookings LRHS Extract File. Standard errors are shown in parentheses.

variable fit significantly better than those with only age in the specification, controlling for time-till-death does not appreciably affect the coefficient on the age variable. This suggests that the relationship between medical costs and age is not a spurious one, due solely to an underlying relationship between health care needs and time until death. It does offer some support, however, for the view that past reductions in mortality rates have led to reduced demand for health care.

We extended our analysis of health status to also consider the effect of the two age measures on retirement decisions. Because the 1969 survey includes a relatively small fraction of retired persons, we report results for both it and the subsequent 1973 survey. The equation we estimate corresponds exactly to that for hospital costs above, except that we replace HCOST with a variable which equals 1 if the person has retired and zero otherwise. The results of estimating the retirement equation are shown in Table 12. As one would expect, both AGE and YTD have significant effects on the probability of retirement. The 1973 estimates imply that an extra year of age increases the probability of retirement by 8.5 percent. Being a year closer to death increases the probability of retirement by 7.2 percent. While these effects appear to be comparable, the important result is again that they are independent. Controlling for time until death does not change the effect of AGE on retirement for either the 1969 or 1973 samples, again suggesting that there is a genuine link between age and retirement. This conclusion is buttressed by the time series evidence. Increases in retirement have occurred rapidly over the past three decades despite major improvements in life expectancy.

Table 12

Reduced Form Retirement Equations, 1969 & 1973 LRHS Samples

	1969 Sample			1973 Sample		
	Equation 1	Equation 2	Equation 3	Equation 1	Equation 2	Equation 3
Constant	-0.724 (0.379)	0.534 (0.059)	-0.363 (0.376)	-4.756 (0.728)	0.919 (0.130)	-4.486 (0.740)
MARRIED	-0.163 (0.034)	-0.155 (0.033)	-0.158 (0.033)	-0.027 (0.061)	-0.011 (0.064)	-0.017 (0.061)
EDUC	-0.017 (0.005)	-0.014 (0.005)	-0.015 (0.005)	-0.030 (0.009)	-0.026 (0.009)	-0.029 (0.009)
SMSA	-0.006 (0.021)	0.008 (0.021)	0.007 (0.021)	0.009 (0.038)	0.020 (0.040)	0.015 (0.038)
RACE	0.107 (0.047)	0.122 (0.046)	0.119 (0.045)	-0.012 (0.079)	-0.006 (0.083)	-0.022 (0.079)
AGE	0.017 (0.006)	---	0.015 (0.006)	0.084 (0.011)	---	0.034 (0.011)
YEARS TO DEATH	---	-0.024 (0.007)	-0.026 (0.007)	---	-0.072 (0.036)	-0.072 (0.036)
SURVIVOR	---	-0.018 (0.042)	-0.008 (0.042)	---	0.099 (0.100)	0.164 (0.096)
SSR	68.9	67.4	66.9	139.7	150.6	138.3
R^2	.059	.093	.099	.096	.026	.106
No. of Obs.	786	786	786	618	618	618

Notes: Standard errors are shown in parentheses. See text for further description.

In results which are not reported here, we found that the YTD variable in both the retirement and hospital cost equations had a highly nonlinear effect. Imminent death has a strong positive effect on the probability of retirement, and is very highly correlated with hospital expenses. Further research could usefully explore these effects in greater detail.

V. Conclusions

Our analysis of several types of data suggests that neither an extremely optimistic nor an extremely pessimistic view of the impact of declining old-age mortality is appropriate. Increased survivorship among relatively unhealthy members of the population has in the past been offset by general reductions in morbidity, leaving the age-specific health status of the population largely unchanged. While projections are difficult, there is no obvious reason to expect this pattern to change.

This suggests that as a reasonable first approximation, future Medicare costs or the costs of institutionalization can be estimated using current age specific information. Projections of this type are somewhat ominous. Based on 1982 age specific rates of institutionalization, and intermediate mortality assumptions, one can forecast a 53 percent increase in the population of institutionalized men, and a 67 percent increase for women by the year 2000. By 2020, the increase will be nearly 120 percent. Under the optimistic mortality assumptions, the corresponding figures are 64 percent for men by 2000, and 79 percent by women. By 2020, both populations have risen to more than 150 percent of their 1982 level. Based on the most recent profile of Medicare costs by age, one can forecast an increase of 37 percent by the year 2000 under the intermediate mortality assumptions. The increase would be 43 percent under optimistic assumptions. Since the intermediate mortality assumption implies that the 65 and over population will increase by 32 percent by the year 2000, these findings suggest that 22 percent of the increase in male institutionalization,

and 35 percent of the female increase, will result from aging of the elderly population. For Medicare, the comparable figures are four and seven percent.

These increases in costs seem large relative to the savings attainable through improvements in the delivery of health care, or the savings that might be possible through making consumers bear a larger fraction of health care costs. This suggests that we must inevitably plan on increases in the share of national income devoted to taking care of the dependent elderly, even if substantial improvements in the delivery of care are achieved.

Proposals to redefine the elderly are frequently advanced. The proponents suggest that the age of 65, originally set by Bismark, is no longer an appropriate demarcation point for defining the elderly. More importantly, it is often argued that with increases in life expectancy, the normal age of retirement should be increased so as to preserve the ratio of working years to retirement years for the average member of the population. Indeed this principle was enshrined in the 1982 Social Security reform package which calls for future increases in the Social Security retirement age. Our analysis suggests that these policy prescriptions are inappropriate. The data do not support Fuchs' view that age should be measured backwards from death. Reductions in mortality do not seem to be associated with reductions in morbidity at each age. There is little reason to think that the health status of the typical 65-year-old twenty years from now will be better than it is now. Hence, there is little basis for proposing a redefinition of the elderly.

Reductions in mortality however will be associated with increases in the variance of health status at any given age. Medical progress will make the

best off members of any given cohort still better off, while marginal survivors are likely to be in very poor health. This suggests the desirability of flexible policies when dealing with the aged population. Policies based on necessarily arbitatry age thresholds will become less and less satisfactory as the dispersion of health status in the population increases.

Our analysis has focused on the effects of mortality reduction taken as a whole, without distinguishing the cause of mortality reductions. Further detailed investigations would be extremely valuable, since different types of reductions in mortality will have different effects. Reductions in accidents, for example, may not change the composition of the population in an unfavorable way, while the implantation of artificial hearts will increase dependency among the surviving population. At the margin, it would probably be desirable to tilt medical progress towards policies of the first type. The criterion of maximizing total years of lifespan extension tends to favor policies directed at persons who, if saved, will be in good health. Maximizing the number of lives saved in the current year, another criterion, is more likely to lead to saving individuals who will be unhealthy and have a high risk of death.

ENDNOTES

1. These data problems are discussed in Ansley Coale and Ellen Kisker, "Mortality Crossovers: Reality or Bad Data?," mimeo, Princeton University, 1985.

2. For a discussion of the evidence on mortality among the extreme aged see Kenneth Manton, "Changing Concepts of Morbidity and Mortality in the Elderly Population," Milbank Memorial Fund Quarterly, 60 (1982), p. 192.

3. U. S. Department of Health and Human Services, Social Security Administration, Office of the Actuary, Social Security Area Population Projections, 1984 Actuarial Study No. 92, May 1984.

4. Cohort life tables were supplied by the Metropolitan Life Insurance Company.

5. We calculate the share of people alive in 1980 who would not have been alive if they faced the m-years earlier life table as $(MS(60,m)*N_{60} + MS(70,m)*N_{70} + MS(80,m)*N_{80} + MS(90,m)*N_{90})/(N_{60} + N_{70} + N_{80} + N_{90})$ where N_i is the number of persons of age i alive in 1980. We use data only on persons every ten years because these are the only cohort life tables available to us.

6. Social Security Administration, Office of the Actuary, op. cit.

7. Health Care Financing Administration, Annual Medicare Program Statistics, 1982.

8. Unpublished data from the 1982 National Master Facility Survey provided by the National Center for Health Statistics.

9. Victor R. Fuchs, "Though Much is Taken: Reflections on Aging, Health, and Medical Care," Milbank Memorial Fund Quarterly 62 (1984), pp. 143-165.

10. James F. Fries, "Aging, Natural Death, and the Compression of Morbidity," New England Journal of Medicine 303 (1980), pp. 130-135; and James F. Fries and L. M. Crapo, Vitality and Aging: Implications of the Rectangular Curve, San Francisco: W. H. Freeman, 1981.

11. Kenneth Manton, op. cit. A similar viewpoint is found in Lois Verbrugge, "Longer Life But Worsening Health?" Milbank Memorial Fund Quarterly 12 (1984), pp. 475.

12. The issues here closely parallel those in the labor economics literature on heterogeneity versus state dependence.

13. James W. Vaupel, Kenneth Manton, and Eric Stallard, "The Impact of Heterogeneity in Individual Frailty on the Dynamics of Mortality," Demography 16 (1979), pp. 439-454.

14. We normalize the average frailty to unity following Vaupel, et. al.

15. We employ data on cohort death rates to study changes in mortality. The analogous correction for these data is:

$$q^j(z_i,t) = 1 - \exp\{-kz_i*[S^j(t+1)^{-1/k} - S^j(t)^{-1/k}]\}.$$

16. Coale and Kisker, op. cit.

17. Some discussion of the importance of heterogeneity in health contexts may be found in Donald Shepard and Richard Zeckhauser, "The Choice of Health Policies with Heterogeneous Populations," in V. Fuchs (ed.), The Economics of Health, University of Chicago Press, 1981.

18. Kenneth Manton, Eric Stallard, and James W. Vaupel, "Methods for Comparing the Mortality Experience of Heterogeneous Populations," Demography 18 (1981), pp. 389-411. The authors also found some cases with much more heterogeneity, corresponding to k of less than 1.

19. We extrapolate observed death rates up to age 72, using an estimated Gompertz function, to obtain estimates of death probabilities at extreme ages.

20. A discussion of the Gompertz function and its uses may be found in S. Horiuchi and A. Coale, "A Simple Equation for Estimating the Expectation of Life at Old Ages," Population Studies 36 (1982), 317-26.

21. Fries and Crapo, op. cit.

22. Fuchs, op. cit.

23. In 1977, Medicaid paid for 51 percent of nursing home residents aged 65-74, 43.9 percent of those between 75 and 84, and 48.6 percent of those over 85. Data are taken from the U.S. Public Health Service, Characteristics of Nursing Home Residents, Health Status, and Care Received, Vital and Health Statistics, Series 13, Number 51, p.41.

24. However, one fifth of all elderly persons will be in a nursing home at some point. See Victor Fuchs, How We Live, Cambridge, Mass., Harvard University Press, 1983.

25. Data are drawn from The Medicare Program Statistics for 1982, and from Health Care Financing: Program Statistics for 1966.

26. This may explain the presence of some outlying observations, such as the one for women aged 75+ in 1965.

27. All estimates were obtained using the Brookings Retirement History Survey extract, provided by Christine de Fontenay. We omitted one observation which claimed that the respondent had died in 1958.

Table A-1

Estimates of Gompertz Models for Mortality Rates, Age 55+

Cohort	α	β	R^2	No. of Observations
1880/Male	-4.103 (0.013)	0.073 (0.0007)	.997	31
1880/Female	-4.531 (0.023)	0.077 (0.0013)	.992	31
1890/Male	-4.169 (0.009)	0.075 (0.0005)	.999	31
1890/Female	-4.786 (0.016)	0.081 (0.0009)	.996	31
1900/Male	-4.247 (0.013)	0.077 (0.0009)	.997	26
1900/Female	-5.017 (0.006)	0.085 (0.0004)	.999	26
1910/Male	-4.215 (0.019)	0.067 (0.0022)	.986	16
1910/Female	-5.001 (0.011)	0.069 (0.0012)	.996	16

Notes: Estimates are based on cohort life table data provided by the Metropolitan Life Insurance Company. The equations estimated are:

$$\log(-\log(1 - q_t^k)) = \alpha^k + \beta^k*(t-55) + \epsilon_t^k.$$

Estimation is by ordinary least squares; standard errors are shown in parentheses.