New Models for Managing Longevity Risk

Public-Private Partnerships

Edited by

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Longevity risk is typically defined as the problem of people living longer than expected. From an aggregate perspective, for example the one adopted by a pension fund manager, increasing life expectancy—that is, an increase in the *average* age at death of a covered population—implies that financial reserves may be inadequate to meet payment obligations. From an individual perspective, the problem is one of 'outliving one's assets,' but that problem can arise from having too few assets as well as from living longer than anticipated. The aggregate form of longevity risk is sometimes characterized as one associated with the *uncertainty* attached to future lifetimes, rather than simply the *length* of future lifetimes (e.g. Brouhns et al. 2002; de Waegenaere et al. 2010). The present chapter is concerned primarily with the length, rather than the variability, of remaining lifetimes.

At its core, longevity risk involves two dimensions: the level of assets, from which an income stream is to be generated, and the length of remaining life, which defines the period of time for which the income stream is needed. This chapter explores the possible role for a third dimension of the problem of longevity risk, namely the role of *active* (or 'disability-free') *life*—an individual-level phenomenon—and *active life expectancy* (ALE), which is an aggregate or cohort-level phenomenon. An individual's 'active' status can change during her or his lifetime, and it can improve as well as worsen (Wolf et al. 2007). Thus, during a person's remaining lifetime, the total period of time spent disabled (or, in the complementary *active* state) can be viewed as a random variable whose values range from zero to the entirety of remaining life.

ALE is defined as the average of the individual-level random variable 'cumulative time spent without disability' (or 'total active life'), and it can refer to either an actual population (i.e. a cohort) or an artificial population (e.g. in period terms). Nevertheless, there appear to be no available data that record the entirety (and few that record even a portion) of the process

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through which time spent with (or without) a disability steadily accumulates during individuals' lives. The most common form of individual-level longitudinal data on disabled or disability-free status presents a series of biannual (or in some cases annual) snapshots of individuals' statuses. An exception is the National Health and Aging Trends Study (NHATS), which includes *monthly* measures of respondents' receipt of help from others with personal-care tasks (such as eating, bathing, and dressing) and mobility-related tasks (such as getting into or out of bed) (Freedman et al. 2015). But to date, at most eight years of these monthly indicators can be produced using available NHATS public-use data, limiting its usefulness in characterizing the full picture of active status from age 65 to death. Instead, a large and vigorous literature has developed to produce estimates of ALE despite an absence of individual-level data on the length of active life, the phenomenon for which ALE is the supposed average (Imai and Soneji 2007; Laditka and Laditka 2009).

In what follows, we first review trends in life expectancy and ALE—aggregate-level phenomena, in both cases—and consider the implications of ALE for longevity risk. The section concludes that ALE trends add little or nothing to the understanding of aggregate longevity risk. The second section of the chapter turns to the individual perspective on longevity risk, considering the implications of active life—or alternatively, life spent with a disability—for longevity risk. Within the constraints imposed by data limitations noted above, we conclude that there are very striking implications of disability status for longevity risk, although given the complexity of the situation revealed by the data, clear-cut behavioral rules based on these findings are not evident. In a brief final section, we discuss interactions between public and private responses to disability late in life.

Trends in Life Expectancy and Active Life Expectancy

It is a well-known fact that period life expectancy has been increasing from year to year—with a handful of exceptions—for many decades. Although the available evidence is more limited, it suggests that ALE has been increasing in recent years as well.

While there are large differences in life expectancy by sex, race, and other background characteristics, for simplicity's sake I will limit attention mainly to both sexes/all races data. Moreover, given the focus of this chapter on longevity risk in retirement, I will limit attention to life expectancy at age 65. Figure 3.1 plots life expectancy as computed by the National Center for Health Statistics for the period 1950 to 2018 (Figure 3.1 also shows estimates of ALE from several sources, as described below). Despite the

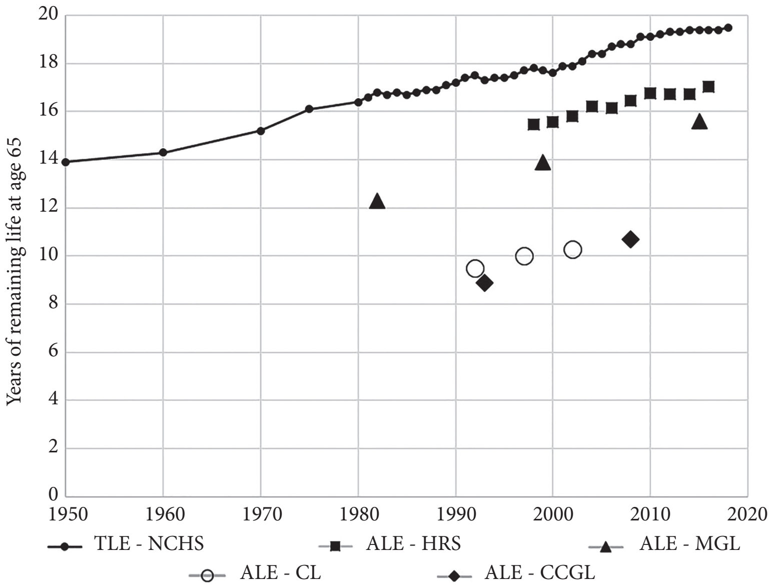

Figure 3.1 Trends in total and active lifetime at age 65

Sources: National Center for Health Statistics (2018) (NCHS); HRS (2019), author's calculations; Manton et al. (2006) (MGL); Cai and Lubitz (2007) (CL); Chernew et al. (2016) (CCGL).

existence of several brief periods of decline, the predominant pattern for life expectancy is clearly one of improvement during this period. A recent analysis of death rates—the inputs into computation of life expectancy—for roughly the same period (1969–2013) found statistically significant 'break points' in the trend line for overall mortality in 1978, 2002, and 2010 (Ma et al. 2015). The downward trend in overall mortality continued, but its slope changed somewhat, at each such breakpoint. An analogous pattern can be seen in the pattern for life expectancy at age 65 in Figure 3.1.

The more-or-less regular trend of increasing period life expectancy over a nearly 70-year period would seem to suggest that pension fund managers could develop forecasting tools that anticipate and therefore eliminate the problem of longevity risk. Moreover, the slowdown—and apparent cessation, since 2014—of the trend toward growing life expectancy at age 65 might, in turn, imply that the problem of longevity risk could diminish on its own.

Turning from life expectancy, defined by the unambiguous and irreversible transition from living to dead, to ALE, defined by a partitioning of remaining life years into those spent with or without disability, introduces several complexities into the analysis. Formally, ALE is the area under a survival curve that is multiplied by the proportion 'active' at each age (Imai and Soneji 2007). In practice, calculating period ALE using the most widely used technique, the so-called 'Sullivan' method, is simple: the calculations entail multiplying elements of the person-years-lived column of a life table by the corresponding elements of an array of age-specific disability (or non-disability) prevalence rates. The great majority of applications of this approach use a binary distinction between 'disabled' and 'disability-free' (or 'healthy' versus 'unhealthy,' or 'active' versus 'inactive,' or any number of other health- or functioning-related categories).

A large literature devoted to ALE has developed since the early contributions of Sullivan (1971) and Katz et al. (1983), and this literature supplies many and varied estimates for ALE. One reason for the proliferation of ALE estimates is the variety of measures of 'active' status that have been used—the distinction between 'disabled' and 'disability-free' is not nearly as straightforward, in concept or measure, as is the distinction between 'alive' and 'dead.' For this chapter, I have produced a series of period ALE estimates for 1998, 2000 . . ., 2016 using disability measures from the US Health and Retirement Study (HRS) in combination with life tables published by the National Center for Health Statistics. The HRS is an ongoing large-scale population-based panel survey that began in 1992, employing biennial surveys thereafter. While the initial sample was limited to the non-institutionalized population age 51–61, various additions to the sample since then permit me to compute ALE estimates for 65-year-olds beginning in 1998 (Health and Retirement Study 2019). The disability measures used here are binary indicators of whether the sample person receives help from another person with any of six Activities of Daily Living (ADLs), namely eating, dressing, toileting, bathing, getting in or out of bed, and walking. Getting help from another person for a basic personal-care task such as these corresponds to a conceptualization of 'disability as dependency'; it is useful in a discussion of longevity risk because help from others entails the use of concrete resources—time or money—and therefore has implications for financial well-being. The ADL indicators used here are taken from a public-use file produced by the RAND Corporation (Bugliari et al. 2019a). The ALE calculations use the abridged life table setup found in Jagger et al. (2014), which implements the Sullivan method for calculating period ALE.

Age-specific disability prevalence rates based on the HRS variables are plotted in Figure 3.2. There are no apparent trends over this 18-year period among the two youngest age groups (65–74 and 80–84), ages at which disability prevalence rates are quite low. Among the older age groups (85–94 and 95+) the trends are not uniform throughout the period, but they are clearly predominantly downward, ending (in 2016) well below their initial values (in 1998).

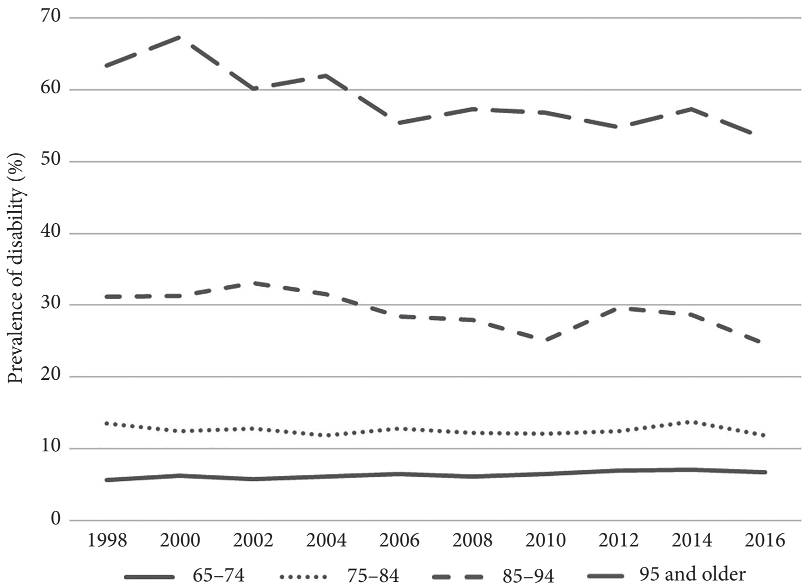

Figure 3.2 Trends in disability prevalence by age group, 1998–2016

Source: Author's calculations based on the Health and Retirement Study (2019).

The 1998–2016 HRS-based calculations of ALE are plotted, along with analogous points taken from three previous publications, in Figure 3.1. All the ALE estimates shown in Figure 3.1 pertain, like the total life-expectancy figures, to the total population; in addition, most also use the Sullivan method. Differences in levels across sources derive principally from the different criteria used to distinguish the 'disabled' from the 'disability-free' portion of the population. Crimmins et al. (2016), for example, define the 'active' population as those free of any limitation of activities, which is the broadest definition used among the sources shown here; this definition, in turn, produces the lowest estimates of ALE. My 'disability as dependency' definition is, in contrast, the most restrictive, and in turn produces the highest estimates of ALE. The other data points plotted in Figure 3.1 use disability criteria that fall between these two extremes. Cai and Lubitz (2007) use the criterion having *difficulty* performing either ADL or Instrumental ADL (e.g. shopping, meal preparation) tasks to define having a disability. Manton et al. (2006) use measures based on the National Long Term Care Survey data, which count as disabled those getting help from other people, along with those using special equipment, to perform daily tasks, and also include those with unmet needs for help dealing with health-related difficulties performing daily tasks.

Figure 3.1 reveals a strikingly consistent pattern of improvement in ALE at age 65 over time. While the levels of ALE differ by the criteria used to measure disability, changes over time in ALE based on each of these measures are close to parallel. Demographers and gerontologists are also interested in the 'compression of morbidity,' a phenomenon associated with the *percentage* of remaining lifetime lived in a disabled state (Cai and Lubitz 2007). With respect to morbidity compression, however, the estimates shown in Figure 3.1 are unclear: when converted to percentages of total life expectancy (not shown), none of the four ALE series plotted in Figure 3.1 shows a clear trend.

Implications for longevity risk

ALE appears to be closely tied to longevity: Figure 3.1 shows that trends in ALE tend to parallel the trend in life expectancy. But whatever the trends in ALE—examined in isolation or relative to total life expectancy—these trends most likely have little or no relevance for aggregate longevity risk. The payout obligations of a pension fund are usually unchanged by the disability status of its beneficiaries. Researchers have pointed out that common factors contribute to improvements in both total life expectancy and ALE (Chernew et al. 2016; Stallard 2016), suggesting that once such factors are incorporated into a mortality forecast, the level or trend in ALE should not add anything to the determination of longevity risk. In other words, whereas ALE is surely associated with the quality of life of pensioners, it seems to have no bearing on the longevity risk faced by pension fund managers.

Active Life and Individual Longevity Risk

In what follows, we discuss individual-level longevity risk from the vantage point of a 65-year-old person, that is, someone close to the typical US age of retirement. At that point, an individual has a given level of assets but faces an uncertain remaining lifetime. If the longevity risk issue is expressed narrowly as the risk of outliving one's assets, then the decision problem facing this individual can be factored into two parts: first, one must—at least implicitly—form an expectation regarding the number of life years that remain; and second, conditional on the first, one must decide either to annuitize assets or to draw them down, presumably following a careful and well-informed plan. The 'annuitize' option, in turn, can be further subdivided into one of buying a fixed-term (usually 10- or 15-year) income stream or buying a lifelong income stream.[1]

With a fixed-term annuity, considerable longevity risk is likely to remain; in contrast, a whole-life annuity eliminates longevity risk but does so at the cost of providing a notably smaller income stream during one's remaining lifetime. Introducing 'active lifetime' as a third factor adds yet another domain of uncertainty into the analysis. A fourth dimension of economic well-being during retirement is, of course, the size of one's social security benefit (for the over 95% of the 65+ population that do, or ultimately will, receive a benefit; Whitman et al. 2011). Although the social security system is subject to political forces and is therefore not immune to adjustment, it is nominally an asset that cannot be outlived; we largely ignore social security in what follows.

We address the dimensions of longevity, wealth, and disability using data from the 1998–2016 HRS, discussed earlier. To the 'disability' indicator already described, we add a measure of the net value of nonhousing wealth, also available in a public-use data file produced by the RAND Corporation (Bugliari et al. 2019b). We also adopt a cohort perspective, focusing on individuals age 65 (or 66) during any of the 1998–2016 interviews, and who were followed thereafter until lost to death or attrition from the sample by the last observation in 2016.[2] This limits the analysis, inasmuch as the maximum age to which surviving members of these biannual cohorts can be followed is 84.

Nevertheless, the most limiting feature of the data used here is the use of biennial indicators of current disability status (i.e. getting help from another person with one or more ADL tasks) as a proxy for continuous but possibly interrupted episodes of disability. At most, we can observe up to 10 biennia—baseline plus follow-up—within the continuously observed cohort subsample, and thereby obtain a partial measure of cumulative (in)active life over the follow-up period. Whether this measure understates or overstates the underlying, but unobserved, continuous active-years measure cannot be determined. If two consecutive HRS interviews are coded 'disabled,' the sample individual cannot be assumed to have been continuously disabled for two years (Wolf and Gill 2008). Similarly, if both of two successive interviews reveal someone to be disability-free at the time of the interview, we cannot assume that the person remained disability-free throughout the interval between interviews.

Moreover, by initiating the observation at age 65, we understate the lifetime experience of disability by an unknown amount: someone coded as disabled at age 65 has been in that state for an unknown period of time, and someone coded as disability-free at age 65 may have had a period with disability that ended at an earlier age. Finally, episodes of disability are right censored by death, given that death and accumulating life years with disability are semi-competing risks: an episode of having a disability can be censored by the event of death, but the reverse is not true (Varadhan et al. 2014). Subject to these data limitations, I report on three relevant random variables: survivorship from age 65, the level of assets held at age 65, and the period of life spent disabled, or free of disability, from age 65 onward. We first consider the marginal distribution of each variable in isolation, and then we consider some relevant associations among them.

Remaining lifetime at age 65

Period-based data on years of remaining life are readily available from conventional life tables. Based on the US period life table for 2007, the midpoint of the years spanned by the HRS sample, life expectancy at age 65 was 18.6 years (Arias 2011). Thus, average remaining lifetime for 65-year-olds slightly exceeded the 18-year follow-up period allowed in the HRS sample used here. Consequently, there is a great deal of censoring of age at death in the cohort sample. Moreover, as is well known, there is a great deal of variability in remaining lifetime (Edwards 2011). For example, the standard deviation of remaining lifetime at age 65, again based on the 2007 period life table for the full US population, was 8.8 years.

The substantial variability in the length of remaining lifetime underscores the challenges people face in deciding how to manage whatever assets they have at the time of retirement. Even if people could do a good job of forecasting the mean of their years-of-remaining-life distribution, there would remain a sizeable probability that they would underestimate the chances of living substantially longer than expected. More problematic, however, is the possibility of bias in people's forecasts of anticipated remaining lifetime. Perozek's (2008) analysis of HRS respondents' answers to survey questions regarding their beliefs about the chances that they will live to age 75, and to age 85, imply survival probabilities that are reasonably accurate for men, but understated for women. Such biases might, in turn, imply that women would tend to draw down their retirement-age assets too quickly. Similar findings appear in McGarry (2022) (Chapter 2, this volume), who additionally demonstrates that HRS respondents tend to adjust downwards their subjective survival probabilities in response to the experience of a serious health shock such as a stroke, the occurrence of heart problems, or receiving a cancer diagnosis.

Assets at age 65

There is a great deal of inequality in wealth holdings in the population generally, and this inequality persists into retirement years (Poterba et al. 2018; Eggelston and Munk 2019). Table 3.1 presents several indicators of the distribution of nonhousing wealth at age 65 in the HRS sample. These indicators are based on a pooled sample of people age 65 or 66 in 1998, 2000. . ., 2016, with wealth values expressed in 2019 dollars. The 2008–2009 years of the global financial crisis are included in this pooled sample, which undoubtedly distorts the wealth picture to some degree. Also, household

TABLE 3.1 Net nonhousing wealth for 65-year-old HRS respondents, 1998–2016

	All	With positive wealth	
Mean	$104,156		$144,489
Median	$6,950		$25,200
20th percentile	$0		$2,880
40th percentile	$1,972		$13,755
60th percentile	$17,850		$44,640
80th percentile	$90,200		$146,300
Given annual rate of return of	2%	4%	6%
percentage with wealth sufficient to buy an annuity providing . . .			
. . . mean annual income . . .			
. . . for 10 years	5.1%	5.6%	6.2%
. . . for 15 years	3.2%	3.8%	4.4%
. . . for life	2.1%	2.7%	3.3%
. . . mean annual income less social security . . .			
. . . for 10 years	11.3%	12.4%	13.3%
. . . for 15 years	8.1%	9.2%	10.4%
. . . for life	5.6%	7.2%	9.6%

Notes:
[a]Annuity pricing data provided by Benny Goodman of TIAA. These calculations are all based on annuity pricing data from pre-Covid-19 crisis times; the public health and financial situations of early 2020 have led to big changes in the demand for, the supply of, and the pricing of retirement annuities.
Source: Author's calculations based on data from Health and Retirement Study (2019).

wealth has been divided by two for married individuals; this may understate people's claims on wealth in the case of an emergency, but it may be reasonably close to what would be available for annuitization.

Confirming results in other studies, Table 3.1 shows that, on average, financial assets at the threshold of retirement are quite low and the distribution of those assets is extremely skewed. Nearly 28 percent of 65-year-olds have negative or zero net assets, and among all 65-year-olds, the average holdings are only $104,156. The median, however, is much lower, $6,950 (the maximum wealth in this sample, not shown in Table 3.1, is over $20 million).

The income streams potentially generated by these wealth holdings would be modest for a great majority of the population. To illustrate this point, Table 3.1 shows the percentage of the age 65 population with wealth sufficient to purchase an annuity that would generate an annual income equal to the population average for this age group, or—more realistically—an average income beyond what is provided by social security on average. For example, Census Bureau data tell us that the average annual income in 2018 of people age 65–74 was $46,325 (US Census Bureau 2019). The average monthly social security benefit among all retired workers age 65–69 in

December 2017 was $1,926 (Social Security Administration 2019), implying an annual benefit of $23,106. Using prototypical annuity pricing, for $100,000 one could purchase a 10-year annuity that, in 2019, generates $917.81 per month, assuming a 2 percent rate of return on the assets used to buy the annuity. Based on these figures, one would have needed $420,601 to buy an annuity sufficient to produce annual income equal to the average annual total income received by individuals in 2018. More realistically, a smaller amount—$193,790—would buy a 10-year annuity that generates income equal to the *difference* between average total income and average social security income, i.e. to 'top off' one's social security benefit. Yet, according to Table 3.1, only 11.3 percent of the age 65 population has enough financial wealth to buy even the smaller annuity. For those with substantial housing wealth and both the ability and willingness to 'downsize,' housing assets could be sold, adding to their annuity purchasing power, but doing so would entail incurring transaction and moving costs. To ensure that one will not outlive one's assets, a whole-life annuity is required, but as Table 3.1 makes clear, only a very small percentage of elders have sufficient assets to achieve an average retirement income using this strategy.

With respect to longevity risk, the main message from Table 3.1—which examines only the asset component of the issue—is that regardless of how long people expect they will live, the great majority of 65-year-olds have either already outlived their assets or they will soon do so. Moreover, any assets they do own will, at best, provide only a modest increment to their social security income. However, the importance of social security to retirees is well documented. For example, one recent study found that roughly half of the aged population lives in households that receive at least 50 percent of their total income from social security, while about one-quarter of the aged live in households that receive at least 90 percent of their family income from social security (Dushi et al. 2017). The data presented in Table 3.1 underscore the fact that the great majority of the population is ill-equipped to achieve a retirement income that is much larger than their social security benefit.

Years of active life

As explained earlier, empirical measurement of years of active life at the individual level is difficult using available survey data. Continuous measures of cumulative time spent disabled do not exist, and the limited time periods covered by most panel surveys place restrictions on the part of the life cycle that can be observed. Using the HRS measures and cohort sample previously described, we can examine the age profile of current and cumulative (since 65) *biennial* indicators of disability. In order to accurately portray the distribution of years of active life beginning at age 65, it would be necessary to model it as a random variable subject to censoring by death; moreover, it would be desirable to build into that model the likely correlation between the two outcomes. That exercise, however, is beyond the scope of this chapter. Instead, I present simple descriptive information on the age profile of disability status among survivors at each age within the age 65 cohort subsample taken from the HRS.

Figure 3.3 shows three age profiles. For each, the data points plotted represent the average in each biennial survey (1998, 2000. . ., 2016) of age at each biennial measurement and (a) the prevalence—that is, the current value—of the disability indicator; (b) the cumulative incidence, during the follow-up period, of the disability indicator; and (c) the total number of biennia for which an individual is coded as 'disabled,' during the follow-up period. These profiles are, of necessity, limited to survivors at each follow-up interview and limited to those for whom complete measurements (from the interview in which they were age 65 or 66) to the present are available. The most that someone can be tracked using this cohort design is 10 waves. Therefore, while the origin of each line includes people age 65 or 66 in all 10 interviews, the second data point is limited to those age 65 or 66 in 1998–2014 and still alive two years later; the tenth and final data point on each line is limited to those age 65 or 66 in 1998 and still alive to respond to the 2016 interview.

Figure 3.3 plots current and cumulative disability patterns from age 65 to roughly the mean age at death—about 84—of survivors, and thus misses the ages where the risk of becoming or remaining disabled are highest (cf. Figure 3.2). The current-prevalence figures rise from about six percent at age 65 to nearly 20 percent at age 84; the cumulative-incidence curve (representing those 'ever disabled' starting with age 65) is only modestly higher, consistent with some degree of recovery among those previously disabled. The average number of biennia with disability curve is well above the other two, but even it suggests only modest 'lifetime' (since age 65) experience with disability among survivors to age 84. Not shown in this figure is the fact that among survivors to the 10th biennial interview, nearly 77 percent have *never* been coded 'disabled.' Thus, the available (and admittedly limited) information presented here suggests that whatever the implications of active life years are for longevity risk, those implications are manifested for only a minority of the population.

The preceding paragraphs have considered each of three dimensions of the longevity risk issue—years of total life, level of assets, and years of active life—in isolation. However, it is likely that these dimensions are associated, so we now turn to a consideration of selected pairwise associations of these dimensions. The analyses undertaken here are intended to be exploratory, not comprehensive.

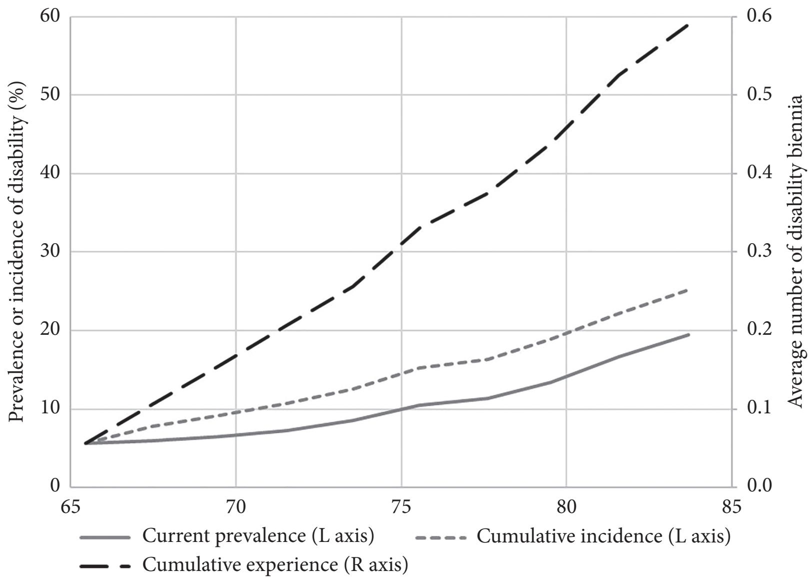

Figure 3.3 Presence of disability by age, HRS respondents, 1998–2016

Note: 'Current prevalence' is the percentage of respondents reporting a disability at the time of interview; 'Cumulative incidence' is the percentage of respondents that have reported a disability at the time of interview or at an earlier interview; 'Cumulative experience' is the average number of interviews, beginning at age 65, at which disability has been reported.
Source: Author's calculations based on data from the Health and Retirement Study (2019).

Wealth predicts longevity

Income has been shown to be strongly associated with longevity (Chetty et al. 2016), and it is to be expected that wealth will have a similar association. To assess the role of wealth at age 65 as a predictor of remaining life years, we estimated a censored regression (i.e. Tobit) model using a pooled sample of individuals observed to be 65 or 66 in 1998. . . 2016 (n = 11,478). As explained before, these individuals are followed for up to nine subsequent biennial interviews, at which time their remaining lifetime is right censored. Others—albeit only a small minority, about 21 percent—are observed to die during the follow-up period. The dependent variable in this regression is the log of survival time. Wealth at age 65 is represented as a categorical variable, with categories corresponding to the quintiles of wealth given that it is positive (those with negative or zero net wealth represented the reference group). The results of this simple bivariate regression are illustrated in Table 3.2. As shown in the table, zero-wealth age 65 individuals are expected to live only 11 more years, while those in each successive quintile of positive net wealth have longer average lifetimes, up to a maximum of 15.5 years for those in the top wealth quintile. Other than for

Table 3.2 Remaining lifetime at age 65, by wealth at age 65

	Zero wealth	Quintile of positive wealth: First	Second	Third	Fourth	Fifth
Average remaining lifetime (years)	11.0	11.6	13.5	14.2	15.1	15.5

Source: Author's calculations based on data from the Health and Retirement Study (2019).

the difference between zero-wealth and first-quintile individuals, all the other longevity differences shown in the table correspond to statistically significant regression coefficients (with $p < 0.0001$).

Wealth is associated with reduced chances of becoming disabled

Just as wealth is positively associated with longevity, it is expected to be associated with a longer life free of disability. To test this hypothesis, we created a pooled sample of person-biennia observations for the HRS cohort sample. Each of the 11,478 baseline individuals used in the preceding analysis are now represented in the analysis sample for as many biennia that they remain alive; the pooled sample thus contains 51,993 observations. The analysis consists of a random-effects logit model of disability prevalence (i.e. 'having a disability' = 1) at each biennial interview, controlling for age, wealth quintile, and the lagged value of 'cumulative number of biennia with disability.' The latter variable is included because when it is equal to zero, the person has not yet experienced any disability, and thus the dependent variable represents the initial onset of disability.

The results of this disability-onset model show the dramatic consequences of wealth as a protective factor against becoming disabled. Figure 3.4 plots the key features of the results: for the ages shown (65–80), the probability of disability onset rises rapidly among those with zero wealth, and much more slowly for those in even the lowest quintile of positive wealth (for the sake of simplicity we show only the first, third, and fifth positive-wealth quintiles). The age profile of disability onset is particularly low for those at the upper end of the wealth distribution: someone previously disability-free at age 80, but with zero wealth at age 65, has a probability of disability onset that is more than 15 times larger (with probability equal to 0.156) than an otherwise comparable person who at age 65 was in the top positive-wealth quintile (with probability of only 0.01).

Disability onset predicts reduced longevity

Finally, we consider the consequences of experiencing disability for one's anticipated remaining lifetime. For this we use the same pooled person-biennium sample just described, but now adopt a random-effects

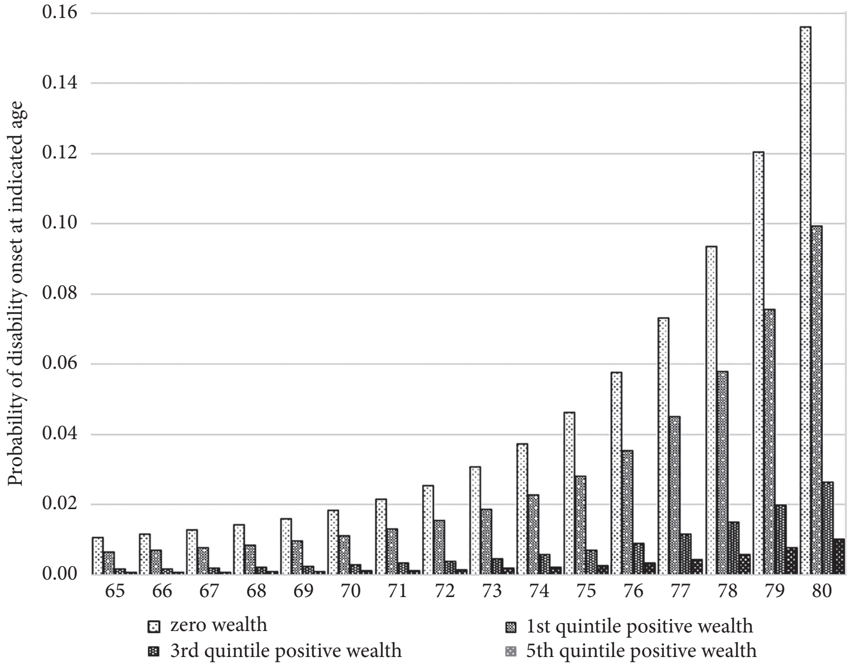

Figure 3.4 Probability of onset of disability by age and wealth at age 65

Note: Heights of bars in this histogram represent the probabilities of first reporting having a disability at each indicated age, for those with zero wealth at age 65 and those in the first, third, and fifth quintiles of positive wealth at age 65.

Source: Author's calculations based on data from the Health and Retirement Study (2019).

Tobit model for remaining years of life at each interview. The controls for disability experience include indicators of one, two, and three or more cumulative biennia with disability; the reference group consists of those with no experience of disability to date. For the reference group, shown in the uppermost line in Figure 3.5, average remaining lifetime is about 15 years at age 65, dropping steadily at later ages (as it must). The most dramatic differences shown in Figure 3.5 are between the disability-free population and those who have, to date, experienced just one biennium of disability (whether current or lagged). For example, a disability-free 65-year-old can expect to live nearly 15 more years, on average; however, someone disabled at that age has a much lower expected remaining lifetime of just nine years. The greater the cumulative experience of disability, the greater the reduction in remaining lifetime, although each additional increment to the number of biennia with disability produces a smaller reduction in remaining lifetime than the one before. Finally, the differences in residual life expectancy between disability classes narrow with age.[3]

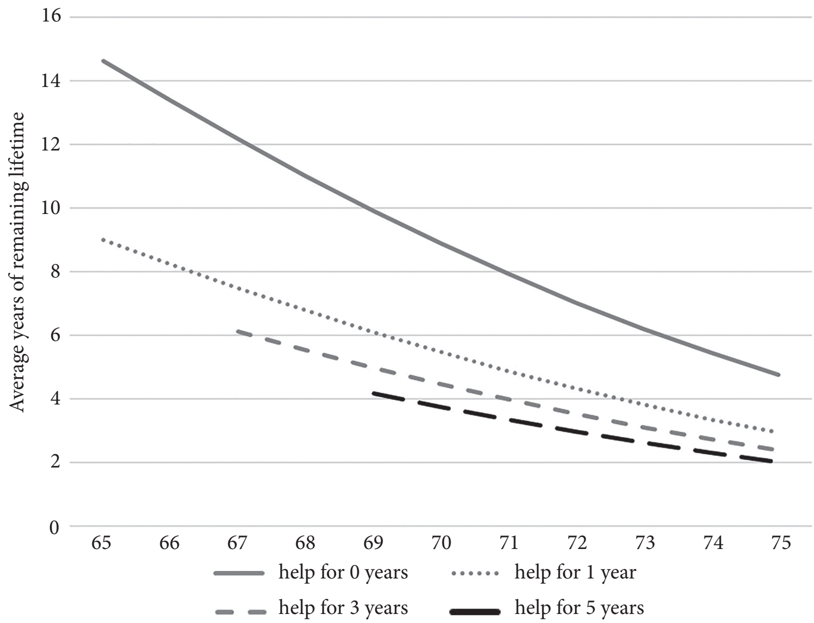

Figure 3.5 Average number of years of remaining lifetime, by current age and duration of current episode of disability

Source: Author's calculations based on data from the Health and Retirement Study (2019).

Longevity Risk, Late-Life Disability, and Public-Private Collaboration

Public-private partnerships, a blanket term that encompasses a broad array of institutional and contractual forms involving a diverse set of public and private actors, have become a major presence in recent decades. It is generally agreed that the sharing of risks across the public and private sectors is a central consideration in these partnerships (Hodge and Greve 2007). With respect to individual-level longevity risk, construed here as outliving one's financial assets, a spectrum of risk sharing arrangements can be observed. A prospective retiree entitled to social security benefits but without any additional private savings or other financial assets faces no longevity risk—presuming the continued operation of the social security program throughout her or his lifetime—thanks to the fact that this risk has been fully transferred to the public sector. Social security is, in effect, the 'public option' for sharing the risk of lack of retirement income. At the other extreme, someone with a pool of financial assets and an intention to self-manage those assets, i.e. to draw them down in keeping with some sort of financial plan, may indeed outlive their private assets, and end up dependent on social security. Private pension plans, and the regulatory apparatus governing them, represent intermediate points along this spectrum of longevity risk.

Introducing active life—or, its complement, late-life disability—into the longevity risk picture adds some interesting features. As shown above, the onset of disability may signal a shorter-than-expected remaining lifetime, which might indicate that assets could be drawn down more rapidly without raising one's longevity risk. On the other hand, the onset of disability signals what might prove eventually to be a need for high-cost long-term care services, which would in turn force the issue of drawing down one's assets more quickly than planned and raise one's longevity risk.

For the disabled older population—defined here as those getting help with everyday tasks—the primary source of help is unpaid family members, or 'informal caregivers.' Johnson and Wiener (2006) using data from 2002 indicate that among those getting help, about 61 percent were community-dwelling individuals getting unpaid help only, while nearly 26 percent were either nursing home residents or community-dwelling individuals receiving paid help only; the remaining 13 percent were community dwellers receiving help from both paid and unpaid sources.

The help provided to disabled elders by family members—mainly by their adult children—might otherwise generate large out-of-pocket costs; family caregiving, in other words, may help avoid longevity risk. It also may delay or completely avert high-cost institutional care (Van Houtven and Norton 2004; Charles and Sevak 2005). By doing so, family caregiving might have as one of its consequences the preservation of bequeathable assets; it might, in other words, have *positive* implications for intergenerational patterns of longevity risk. Yet family caregiving is widely understood to impose substantial costs on the individuals that provide it, and on the families of which they are a part. For example, according to a 2014 survey about 60 percent of informal caregivers are, or at one time while providing care were, employed (NAC/AARP 2015). Alternatively—changing denominators—about 18 percent of the employed population are simultaneously engaged in care provision (Wolf 2019). Employed caregivers report a broad range of care-related costs, among which are missing work, taking unpaid leave (and consequently lowering income), changing to a lower paid or part-time job, or quitting work entirely (Witters 2011; NAC/AARP 2015). Thus, there are good reasons to imagine that family caregiving has *negative* implications for intergenerational patterns of longevity risk.

For those disabled elders whose care is provided in the community by paid caregivers, or in nursing homes (and, therefore, necessarily by paid caregivers), the majority of care costs are borne by Medicaid (Reaves and Musumeci 2015). Paid care services are strongly connected to longevity risk, in view of the fact that eligibility for Medicaid depends in part on passing a stringent asset test. Moreover, there is a great deal of state-to-state variation in the stringency of the asset test: in 2018, for example, the level of allowable assets ranged from a low of $1,500 (in New Hampshire), to a high of $7,560 (in California and South Carolina) (Musumeci and Chidambaram 2019).[4] Thus, in order to be Medicaid eligible, someone must have either arrived at old age without assets or have 'spent down' their assets—quite likely on paid care services—to the low level needed to achieve eligibility. Moving into a nursing home as a self-pay patient is one way of accomplishing spend-down. One analysis of survey data linked to administrative data showed that, among community residents not initially enrolled in Medicaid, fewer than three percent were observed to enroll in Medicaid during a four-year follow-up period if they never moved into a nursing home; meanwhile, 22 percent (20 percent, 19 percent, 17 percent) did transition onto Medicaid by the end of the follow-up period if they had moved into a nursing home within one year (or 2, 3, or 4 years, respectively) after baseline (Spillman and Waidmann 2014). The high cost of nursing homes, in other words, provides the means by which many people deplete their financial assets.

Together, these facts support a somewhat oversimplified characterization of one form of public-private 'partnership' that has arisen to cope with the intersection of longevity risk with active life, as follows: the **private** component of risk sharing consists of family caregivers providing needed care services, protecting insofar as possible their parents' assets (and, indirectly, their own inheritances), while saving the public the expense of paying for 'formal' care services. Under this approach, the costs that do arise are borne narrowly by individuals and their families and may even have intergenerational repercussions. The **public** part of this partnership consists of the provision of expensive care services, in the community or in a nursing home, to a population of disabled elders that have been impoverished by the process of establishing Medicaid eligibility. Those in the latter group have necessarily outlived whatever assets they once had. Costs are borne broadly by society through the taxpayer-funded Medicaid program. This is, at best, a rather haphazard 'partnership' (hence the quotation marks around that word), is one that is not the result of a deliberate design, and is moreover one with substantial between- (and probably within-) state variability.

There is limited space in the aforementioned partnership for private institutions other than families, for example, private long-term care insurance. Indeed, 30 years ago Pauly (1990) showed how nonpurchase of private insurance was a rational response to (among other things) the availability of family members as potential providers of care. Rates of private coverage of this risk continue to be quite low (Brown et al. 2012). Nevertheless, we will mention two policy domains that have the potential to alter the terms of this public-private partnership.

Consumer-directed care

Consumer-directed care refers to an increased role of consumers in managing their own health and health care services. Consumer direction is believed to improve the quality while lowering the costs of health care (Buntin et al. 2006). In the area of long-term care, the principal manifestation of this policy idea has been the Cash and Counseling program, which has been implemented on a limited basis in 15 states (De Milto 2015). In this program, Medicaid beneficiaries eligible for long-term care services receive a cash budget with which to purchase services, and the cash can be used to pay family members to provide at-home care. Because this program operates as a component of Medicaid, the asset-depletion feature of Medicaid is in force here, as well. Yet, by allowing payments to family members, Cash and Counseling could offer a means of reducing some of the costs that would otherwise be borne by family caregivers.

Paid leave

Paid family caregiving leave provisions may allow some people to receive 'paid care' from an employed family member while not requiring that the benefit—the payment for care services—be conditioned on the depletion of the care recipient's assets—i.e. on Medicaid eligibility. While this type of benefit is both capped and limited in duration, it (like Cash and Counseling benefits) has at least some potential to offset what would otherwise be a cost borne by family caregivers; consequently paid caregiving leave has the potential to change the nature of the public-private partnership that has arisen with respect to elder care.

As of 2019, eight states plus the District of Columbia had passed laws mandating the provision of some form of paid family leave by nearly all private employers (National Partnership for Women and Families 2019). Because some of these states have relatively large populations, in 2018 nearly 21 percent of total US employment was in the four states that by then had implemented a paid family leave law; over 28 percent of total US employment (in 2018) was in states that by 2023 will have implemented their paid family leave program.[5] In all cases, these laws require that paid leave be extended to those providing care to parents as well as to the parents of newborn children, the more typical target of such policies.

The first Federal requirement for paid care-related leave emerged with the Families First Coronavirus Response Act (FFCRA) and in the Coronavirus Aid, Relief, and Economic Security (CARES) Act, both of which were passed in March 2020. Both bills were enacted in response to the major and rapidly evolving public health and economic crises associated with the COVID-19 pandemic of 2020. Prior to 2020, Federal policy governing family caregiving leave for private-sector workers was limited to the Family and Medical Leave Act of 1993, which mandated the provision of unpaid leave while exempting those working for small firms and those who fail to meet length-of-service requirements (Klerman et al. 2014). While the FFCRA and CARES Act represent a major shift in Federal leave policy, neither is likely—nor are they intended—to alter the extent or costs of family-provided care to disabled elders: benefits in the new programs are closely tied to quarantining associated with COVID-19 exposure, or to the needs of children unable to attend school. Moreover, the programs were temporary, expiring on December 31, 2020. These initiatives may, however, prove to be the first step toward a more permanent and widespread Federal paid-leave mandate.

Conclusion

Longevity risk can be analyzed as either an aggregate or an individual-level issue. From the aggregate perspective—for example, that adopted by a pension fund manager—an imbalance between the average lifetimes of those in a covered population and the adequacy of fund reserves appears to be little altered by consideration of ALE for that population. Individual-level longevity risk—the prospects for outliving one's financial assets—is, as well, of little relevance to the substantial proportion of the population that reaches retirement age with zero or only modest asset levels. For those people, the problem is one of living on one's social security benefit rather than worrying about annuitizing or drawing down one's savings. Nevertheless, there are important associations among the three dimensions of remaining lifetime, one's level of financial assets, and the experience of disability. The onset of disability in late life provides a signal about the length of one's active life (the individual-level variable of which ALE is a population average); this signal, in turn, indicates that remaining lifetime will be shorter than expected, or that care costs will be greater than expected, or both.

Care needs for most people are addressed through the provision of unpaid care services provided by family members, and this type of care may preserve one's assets as well as protect one's children's inheritance, albeit at what might be a substantial cost to the care providers. In the absence of informal care from family members, or the presence of care needs too severe to be met by family members, one's care needs may end up being publicly financed through Medicaid, but this outcome will generally be accompanied by the exhaustion of one's assets—a full-scale realization, in other words, of longevity risk.

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Julia Carboni and Benny Goodman made helpful contributions to this chapter.

Notes

1. The annuities market offers a number of additional features and variations on these basic plans, which for the sake of simplicity I ignore.
2. The two-year sample inclusion is a consequence of the biennial interviewing design used by the HRS. Thus, respondents age 65 or 66 in 1998 are 67 or 68 in 2000, 69 or 70 in 2002, and so on. The same approach is applied to those age 65 or 66 in 2000, 2002, or later, but with correspondingly shorter follow-up periods.
3. The results plotted in Figure 3.5 are admittedly, but to an unknown extent, an artifact of the functional form and model specification adopted for this purely descriptive analysis. I have not, for example, explored interactions between age and cumulative experience of disability, both of which are time-varying covariates.
4. Arizona was the only state not to impose an asset test on seniors and people with disabilities in 2018.
5. Author's calculations using the annual Local Area Unemployment Statistics data from US Bureau of Labor Statistics (2020).

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