Perfecting Data
HISTORICAL METHODS, April–June 2010, Volume 43, Number 2

Routledge
Taylor & Francis Group

Decennial Life Tables for the White Population of the United States, 1790–1900

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Abstract. In this article, the author constructs new life tables for the white population of the United States in each decade between 1790 and 1900. Drawing from several recent studies, he suggests best estimates of life expectancy at age 20 for each decade. These estimates are fitted to new standards derived from the 1900–1902 rural and 1900–1902 overall death registration area life tables using a two-parameter logit model with fixed slope. The resulting decennial life tables more accurately represent sex- and age-specific mortality rates while capturing known mortality trends.

Keywords: demographic history, demography, life table, mortality, nineteenth century, United States

Life tables summarize the effects of age-specific mortality rates on a real or synthetic cohort. In addition to their descriptive value, life tables are an indispensable tool for demographers, with many applications in the study of mortality, fertility, migration, and population growth. Life tables are often used in conjunction with indirect estimation methods for the study of populations covered by a census but lacking a reliable vital registration system, such as many populations in developing countries and populations in the past. Life tables, for example, can be used to estimate vital rates from census age distributions and are required in own-child fertility analysis (see United Nations 1983, for a description of commonly-used indirect methods).

Demographic historians of the nineteenth-century United States depend heavily on life tables and indirect estimation methods. Although the federal government conducted a census every ten years, it did not implement a vital registration system until the start of the twentieth century (the system was not complete until 1933).[1] As a result, the timing and contours of the demographic transition in the United States are less precisely known than that in nations such as England and Australia, which had comprehensive birth and death registration by the mid–nineteenth century (Jones 1971; Woods 2000). Despite this limitation, demographic historians have been able to estimate annual and age-specific birth rates, net migration rates, and cohort trends in life cycle experiences as far back as the early nineteenth century using census data, life tables, and indirect methods (Yasuba 1962; Coale and Zelnik 1963; Kuznets 1965; Uhlenberg 1969; McClelland and Zeckhauser 1982; Tolnay, Graham, and Guest 1982; Ferrie 1996; Hacker n.d.)

Unfortunately, the results of these studies depend on a small number of life tables, which suffer from limited geographic coverage, limited temporal coverage, and a variety of source-based problems. The earliest life tables rely heavily on data from Massachusetts, a small state in the Northeast characterized by much higher levels of urbanization, industrialization, and immigration and much lower levels of nuptiality and fertility than the nation as a whole. Given the high short-term variability in mortality rates that was characteristic of the nineteenth-century United States, it is also unclear whether life tables based on a single year of data can be used to represent mortality in a year other than the one for which it was constructed. The failure of existing life tables to capture suspected long-term trends in mortality is perhaps their most significant limitation. With just a handful of life tables covering the entire nineteenth century, researchers have been forced to make crude assumptions about long-term mortality trends to conduct their analyses. As discussed in more detail below, recent research indicates that earlier assumptions of long-term mortality decline are in error. Mortality increased significantly in the mid-nineteenth-century United States before beginning its long-term decline.

This article constructs new life tables for the white population of the United States in each decade between 1790 and 1900. The first part of the article reviews research on

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the level and trend of nineteenth-century mortality. Drawing from several recent studies, it suggests best estimates of male life expectancy at age 20 for each decade. The second part of the article investigates sex differentials in mortality and suggests best estimates of female life expectancy at age 20 for each decade. The third part of the article reviews research on the age pattern of male and female mortality. The results indicate that age-specific mortality rates in the nineteenth century did not match the two most frequently cited standards: the west model of the Princeton regional model (Coale, Demeny, and Vaughan 1983) life tables or the 1900–1902 Death Registration Area (DRA) life tables for the United States. It concludes, however, that the life tables constructed for the *rural* part of the 1900–1902 DRA likely approximate the age pattern of nineteenth-century mortality. Finally, the fourth part of the article fits the decennial estimates of life expectancy at age 20 to new standards derived from 1900–1902 rural and overall life tables using a two-parameter logit model with fixed slope. The resulting decennial life tables, it is argued, more accurately represent sex- and age-specific mortality rates while capturing known mortality trends.

The Level and Trend in Nineteenth-Century Mortality

Table 1 shows life expectancy and infant mortality estimates from selected United States life tables in the period between 1789 and 1902 by year of publication (for a more complete listing, including life table summaries for selected cities, see Haines 1998). The tables were constructed from a wide variety of sources, including local bills of mortality, state and national death registration data, census data, family genealogies, and biographical data on special populations such as legislators and college graduates. Table 1 also shows the sex mortality differential at age 20, defined as the female life expectancy at age 20 minus male life expectancy at age 20.

Edward Wigglesworth (1793) constructed the first United States life table using Bills of Mortality for 35 New England towns in the late eighteenth century. Ezekiel B. Elliott (1858), John S. Billings (1885), Samuel W. Abbott (1898), and James W. Glover (1921) relied on death registration data in Massachusetts—the first state to implement a death registration system—to calculate life tables for selected years in the late nineteenth century.[2] Levi Meech (1898) constructed the first national life table. The lack of national death registration data forced Meech to rely on an indirect approach. He estimated cohort declines from the 1830–60 federal censuses, made adjustments from immigration data to account for the lack of a closed population and used retrospective mortality data published by the 1860 census to establish the age pattern of death (1898, 255–59). The creation of a national death registration area (DRA) in 1900 greatly facilitated the creation of life tables. Glover's (1921) 1900–1902 DRA life tables relied on registration data from the 10 states and the District of Columbia that comprised the nation's original DRA. These tables have been widely used by researchers to represent the level and age pattern of mortality in the United States at the turn of the twentieth century.

Two studies conducted in the mid-twentieth century have been widely cited as representative of nineteenth-century mortality. Abram J. Jaffe and William I. Lourie Jr. (1942) relied on death registration data collected by 44 New England towns, several midsized cities, and a few larger cities to construct life tables for the period 1826–35. The results indicated large differentials in life expectancy between rural areas and large urban centers, with life expectancy at birth almost 15 years higher in the selected towns than in the large cities of Boston, New York, and Philadelphia. Paul H. Jacobson's (1957) 1849–50 life tables were based on retrospective mortality data collected by the 1850 census. Jacobson confined his analysis to data collected for Massachusetts and Maryland, reasoning that an arithmetic mean of their age-specific death rates would approximate those for the nation as a whole.

The life table estimates in table 1 are sorted by year of publication to emphasize our relatively recent knowledge about nineteenth-century mortality. Researchers requiring life table data prior to the late 1970s were limited to a handful of tables, which led to great uncertainty about mortality trends. Inferring mortality trends in the early nineteenth century from existing life tables was especially problematic. Warren S. Thompson and Pascal K. Whelpton (1933, 230–31) calculated a slow decline in the crude death rate from 27.8 per 1,000 in the late eighteenth century to 21.4 per 1,000 in 1855 by interpolating between the Wigglesworth (1793) and Elliott (1858) life tables. Reviewing the more recent evidence available to them in the late 1950s, Conrad Taeuber and Irene B. Taeuber (1958, 269) found no conclusive evidence of mortality decline in the first half of the nineteenth century. Yasukichi Yasuba (1962, chap. 3) saw evidence of mortality increase in the few decades preceding 1860 associated with increasing urbanization and declining sanitary conditions. Richard Easterlin (1977), in contrast, argued that increasing per capita income more than offset the negative impact of urbanization and cited life expectancy estimates from the Wigglesworth (1793) and Jacobson (1957) life tables as evidence of significant mortality decline.

Most early observers agreed that the latter half of the century was characterized by substantial mortality decline; although opinions differed about the date of its onset. Taeuber and Taeuber (1958, 269) thought the evidence suggested an "almost continuous" decline in mortality beginning about 1850. To conduct their classic study of long-term trends in white birth rates, Ansley Coale and Melvin Zelnik (1963) assumed a linear decline in mortality between Jacobson's (1957) 1849–50 life tables and the 1900–1902 DRA life tables. In separate analyses based on Simon Kuznets's (1965) census-survival estimates of crude death rates, however, Edward Meeker (1972) dated mortality decline after 1880, when the public health and sanitation movement became more

TABLE 1. Life Tables Estimates of Infant Mortality Rate (IMR) and Life Expectancy (*e*) at Selected Ages, United States, 1798–1901

Investigator(s)	Date of publication	Population	Period	Male IMR	Male e_0	Male e_{20}	Female IMR	Female e_0	Female e_{20}	SMD_{20}
Wigglesworth	1793	Selected Massachusetts towns	1789		36.5 [a]			36.5 [a]		—
Elliott	1858	166 Massachusetts towns	1855	0.155 [b]	39.8 [b]	39.9 [b]	0.155 [b]	39.8 [b]	39.9 [b]	
Billings	1886	Massachusetts	1878–82	0.181	41.7	42.2	0.153	43.5	42.8	0.6
Billings	1886	New Jersey	1879–80	0.152	45.6	43.3	0.131	48.0	44.5	1.2
Meech	1898	U.S. Whites	1830–60	0.162	41.0	40.9	0.134	42.9	41.4	0.5
Abbott	1899	Massachusetts	1893–97	0.172	44.1	41.2	0.147	46.6	42.8	1.6
Glover	1921	Massachusetts	1890	0.168	42.5	40.7	0.148	44.5	42.0	1.3
		DRA states	1900–02	0.133	48.1	42.2	0.111	50.9	43.7	1.6
		DRA rural areas	1900–02	0.109	54.0	45.9	0.090	55.4	46.0	0.1
Jaffe and Lourie	1942	44 New England towns	1826–35			42.9 [b]			42.9 [b]	
		Salem, MA, New Haven, CT	1826–36			37.8 [b]			37.8 [b]	
		Boston, New York City, Philadelphia	1826–37			28.0 [b]			28.0 [b]	
		Estimated United States				41.7 [b]			41.7 [b]	
Jacobson	1957	Massachusetts and Maryland whites	1849–50	0.161	40.4	40.1	0.131	43.0	41.7	1.6
Vinovskis	1972	Massachusetts	1859–61		46.4	44.0		47.3	43.0	−1.0
Kunze	1979	Genealogies	1800–14	—		44.7	—		43.4	−1.3
			1815–29	—		44.1	—		43.3	−0.8
			1830–34	—		44.4	—		47.8	3.4
			1835–39	—		44.6	—		42.3	−2.3
			1840–44	—		43.5	—		41.7	−1.8
			1845–49	—		42.5	—		40.7	−1.8
			1850–54	—		42.2	—		40.1	−2.1
			1855–59	—		40.8	—		40.5	−0.3
			1860–64	—		40.5	—		40.2	−0.3
			1865–69	—		40.1	—		40.1	0.0
			1870–74	—		42.6	—		40.0	−2.6
			1875–79	—		43.7	—		41.5	−2.2
			1880–84	—		44.4	—		41.8	−2.6
			1885–89	—		43.8	—		42.2	−1.6
Pope	1992	Family Histories	1780–99	—		47.4	—		45.6	−1.8
			1800–09	—		46.4	—		47.9	1.5
			1810–19	—		44.6	—		44.4	−0.2
			1820–29	—		43.3	—		44.9	1.6
			1830–39	—		44.6	—		44.6	0.0
			1840–49	—		41.5	—		37.1	−4.4
			1850–59	—		40.8	—		39.5	−1.3
			1860–69	—		41.2	—		42.2	1.0
			1870–79	—		44.3	—		42.2	−2.1
			1880–89	—		45.8	—		42.9	−2.9

(Continued on next page)

TABLE 1. Life Tables Estimates of Infant Mortality Rate (IMR) and Life Expectancy (*e*) at Selected Ages, United States, 1798–1901 *(Continued)*

Investigator(s)	Date of publication	Population	Period	Male IMR	Male e_0	Male e_{20}	Female IMR	Female e_0	Female e_{20}	SMD_{20}
Kasakoff and Adams	1995	New England families	1750–59[c]			48.1				
		(settled before 1650)	1760–69[c]			48.3				
			1770–79[c]			43.2				
			1780–89[c]			44.5				
			1790–99[c]			43.6				
			1800–09[c]			42.0				
			1810–19[c]			40.7				
			1820–29[c]			40.6				
			1830–39[c]			39.0				
Ferrie	1996	Native-born whites (weighted)	1850			45.4				
		Urban (weighted)	1850			38.0				
		Rural (weighted)	1850			47.6				
		Foreign-born whites (weighted)	1850			35.7				
		Urban (weighted)	1850			30.6				
		Rural (weighted)	1850			45.3				
Hacker	1996	Yale graduates	1790–1829			40.1				
Levy	1996	Maryland legislators	1750–1764[c]			34.7[d]				
		South Carolina legislators	1750–1764[c]			31.7[d]				
Haines	1998	U.S. white population (U.S. model)	1850	0.228	38.4	38.8	0.206	40.6	40.2	1.4
		U.S. white population (West model)	1850	0.195	38.8	37.5	0.155	43.5	40.8	3.3
		U.S. white population (U.S. model)	1860	0.188	43.2	41.0	0.175	44.1	41.7	0.7
		U.S. white population (West model)	1860	0.165	43.0	39.6	0.139	46.2	42.1	2.5
		U.S. white population (U.S. model)	1870	0.185	44.1	41.7	0.166	46.4	43.3	1.5
		U.S. white population (West model)	1870	0.156	44.4	40.2	0.126	48.5	43.2	3.0
		U.S. white population (U.S. model)	1880	0.214	40.4	40.0	0.215	40.6	40.9	0.9
		U.S. white population (West model)	1880	0.180	40.9	38.5	0.154	43.8	40.9	2.4
		U.S. white population (U.S. model)	1890	0.157	46.0	41.7	0.145	47.4	42.8	1.1
		U.S. white population (West model)	1890	0.148	45.6	40.8	0.124	48.9	43.4	2.6
		U.S. white population (U.S. model)	1900	0.128	48.5	42.0	0.112	50.7	43.5	1.5
		U.S. white population (West model)	1900	0.135	47.8	41.7	0.109	51.7	44.6	2.9
Hacker[e]	n.d.	Princeton graduates	1790–1819			36.2				

Note. SMD = sex mortality differential at age 20, defined as the female life expectancy at age 20 minus male life expectancy at age 20; DRA = death registration area.
[a]Both sexes combined. Vinovskis revised estimate.
[b]Both sexes combined.
[c]Birth cohorts.
[d]e_{20} estimated from e_{25} (Maryland legislators = 32.2; South Carolina legislators = 29.6).
[e]Estimate made by author from biographical information published in McLachlan 1976; Harrison, McLachlan, and Woodward 1980; Harrison 1981; Woodward and Craven 1991; and Looney and Woodward 1991.

effective, whereas Robert Higgs (1973) observed a decline in rural areas from the 1870s.

Beginning in the 1970s, new research considerably clarified our understanding of nineteenth-century mortality. Much of the new research was critical of earlier studies. In a series of articles, Maris Vinovskis (1971, 1972, 1978) evaluated the Wigglesworth (1793), Jaffe and Lourie (1942), Elliott (1858), and Jacobson (1957) life tables, all of which relied on data from Massachusetts. Although the Wigglesworth life table suggested a reasonable estimate of life expectancy at birth, Vinovskis (1971, 589) observed that Wigglesworth (1793) lacked adequate data on the age distribution of the towns, which required adjustments amounting to "little more than intelligent guessing." Vinovskis also noted that the towns covered by the Wigglesworth life table were not representative of other New England towns in important characteristics, including their relative affluence and degree of urbanization, making it difficult to evaluate the table's representativeness. Vinovskis (1972, 204–5) faulted Jaffe and Lourie (1942) for relying on data from many small towns with under-registered deaths, thus overestimating the significance of the rural-urban differential in mortality and understating the overall level of morality. Elliott (1858), Vinovskis (1972, 208–10) argued, erred in the opposite direction. To avoid including places with deficient record keeping, Elliott eliminated towns with a crude death rate of less than 16 per 1,000. In doing so, however, Elliott likely removed towns whose true death rate was lower than 16 per 1,000 and thus overstated the true level of mortality. Vinovskis (ibid.) also contended that Elliott's reliance on just one year of mortality data was problematic, given the era's high short-term variability in death rates. Finally, Vinovskis demonstrated that Jacobson (1957) failed to consider contemporary critiques of the 1850 census of mortality, which noted that deaths were unevenly registered, and failed to consider that the census was taken during a cholera epidemic, resulting in a likely overestimation of mortality despite the under-registration of deaths. Given these critiques, it is no surprise that Vinovskis (1978) strongly cautioned against inferring mortality trends from the Wigglesworth (1793), Jaffe and Lourie (1942), Elliott (1858), and Jacobson (1957) life tables. Drawing from bills of mortality and state registration reports, Vinovskis concluded that there was little trend in Massachusetts mortality during the first half of the century.

Meech's (1898) life table also received an extensive critique. In a detailed reconstruction and analysis, Michael R. Haines and Roger C. Avery (1980) noted that Meech was forced to make a number of assumptions to construct his life table, including the questionable assumptions that the underenumeration of deaths in the census and the required adjustment of gross to net migration were independent of age. As a result, Haines and Avery concluded that the Meech life table likely underestimated infant mortality and overestimated early childhood mortality; although it gave reasonable results overall.

Finally, a number of researchers have cautioned against inferring national mortality patterns from life tables constructed for Massachusetts and the 1900–1902 DRA (Easterlin 1977, 133; Condran and Crimmins 1979, 1; Preston and Haines 1991, 49–50; Haines and Preston 1997). Although these tables were based on relatively well-reported death registration data[3] and are thus reasonably accurate descriptions of the level of mortality and sex- and age-specific mortality patterns in those areas—they are unlikely to be representative of the national population. Table 2 compares the population of Massachusetts, the 1900–1902 DRA,

TABLE 2. Comparison of Selected Characteristics of Massachusetts and the Original Death Registration Area (DRA) States with the United States as a Whole, 1850 and 1900

	1850		1900			
	U.S.	Massachusetts	U.S.	DRA	DRA—"rural"	Massachusetts
Total population	100.0	5.3	100.0	26.2	12.0	3.7
Percentage urban	16.9	51.6	38.8	60.1	13.2	67.9
Percentage foreign born	11.3	18.0	13.8	22.6	14.0	29.7
Labor force						
Percentage in agriculture	50.5	22.8	35.5	17.6	39.0	5.2
Females age 20–49						
Average number of own children in household	2.29	1.46	1.87	1.49	1.63	1.31

Notes. The original death registration states of 1900 consisted of the six New England states (Connecticut, Maine, Massachusetts, New Hampshire, Rhode Island, and Vermont), Indiana, Michigan, New Jersey, New York, and the District of Columbia. Rural areas of the DRA were initially defined as places with less than 8,000 inhabitants. Later census definitions changed the urban/rural threshold to places of 2,500 inhabitants.

Source: 1850 and 1900 IPUMS samples from Steven Ruggles, Matthew Sobek, J. Trent Alexander, Catherine A. Fitch, Ronald Goeken, Patricia Kelly et al., *Integrated Public Use Microdata Series: Version 4.0* (Minneapolis, MN: Minnesota Population Center, 2009) [producer and distributor].

and the overall United States in 1850 and 1900 using data from the 1850 and 1900 Integrated Public Use Microdata Series (IPUMS) samples (Ruggles, Sobek, Alexander, Fitch, Goeken, Kelly et al. 2009). Massachusetts was much more urban than the rest of the nation, had a proportionately larger and more rapidly growing foreign-born population and had a much lower proportion of its labor force engaged in agriculture (the state was one of the first to industrialize in the early nineteenth century). Moreover, Massachusetts enjoyed one of the best public health systems in the nation and was the leading state in the employment of women in the labor force and in the fertility transition. Massachusetts women age 20–49, for example, had an average of just 1.5 co-residing own-children in 1850 and 1.3 in 1900, suggesting fertility rates approximately one-third lower than that of the nation as a whole. Thus, although Massachusetts has the best available mortality data for the nineteenth century, its level, trend, and age pattern of mortality are unlikely to be representative of the United States as a whole.

Table 2 also indicates that the population of the 1900–1902 DRA was not representative of the nation. The initial DRA included the six New England states, New York, New Jersey, Michigan, Indiana, and the District of Columbia. Although the 1900–1902 DRA was much larger than the state of Massachusetts—representing about 26.2 percent of the nation's population in 1900 compared with only 3.7 percent for Massachusetts—it varied from the rest of the nation in similar, if less dramatic, ways. The DRA was more urban than the United States as a whole and its population included a higher proportion of foreign-born residents and a lower proportion of agricultural workers. Women in the DRA had an average of 20 percent fewer co-residing children in the household than women in the nation as a whole. It is noteworthy, however, that differences between the *rural* parts of the 1900–1902 DRA and the rest of the nation were less extreme. Rural parts of the DRA included about the same proportion of foreign-born residents and workers engaged in agriculture. Fertility rates in rural areas of the DRA were much closer to the national average.

Fortunately, just as confidence in the representativeness and accuracy of existing life tables was falling, new research significantly enhanced our understanding of nineteenth-century mortality trends. Beginning with Michael Haines's analysis of the United States censuses of mortality (1979) and Kent Kunze's (1979) and Robert W. Fogel's (1986) demographic analyses of family genealogies, life expectancy estimates have accumulated for each decade of the nineteenth century. Clayne Pope's (1992) study of family histories is perhaps the most significant contribution for the first half of the century while Haines's (1998) construction of life tables for the white, black, and overall populations is the most important work for the last half of the century; although important research has also been published by Richard S. Meindl and Alan C. Swedlund (1977); Gretchen A. Condran and Eileen Crimmins (1979, 1980); Crimmins (1980); Daniel Scott Smith (1982, 2003); Condran and Rose A. Cheney (1982); Cheney (1984); Stephen Kunitz (1984); Condran (1987); Richard Steckel (1988); Barbara J. Logue (1991); Eric Leif Davin (1993); Alice Kasakoff and John Adams (1995, 2000); Joseph Ferrie (1996, 2003); Antonio McDaniel and Carlos Grushka (1995); J. David Hacker (1997); John E. Murray (1997, 2000); Chulhee Lee (1997, 2003); Susan I. Hautaniemi, Alan C. Swedlund, and Douglas L. Anderton (1999); Anderton and Susan Hautaniemi Leonard (2004); and Jeffrey K. Beemer, Anderton, and Hautaniemi Leonard (2005).

Several of the newer studies—including those by Haines (1979, 1998), Ferrie (1996, 2003), and Condran and Crimmins (1979, 1980)—have relied on retrospective mortality data collected by the Census Office/Bureau of the Census between 1850 and 1900. Beginning In 1850, census marshals were instructed to record the name of every person in the household who died in the year prior to the census, as well as the person's age, sex, race, marital status, occupation, and cause of death. The collected data were tabulated and published in separate mortality volumes. These tabulated data appear to be tailor-made for the construction of life tables: the number of deaths at each age and sex can be used as the numerator in the calculation of age-specific death rates while the denominator for the midyear population in each age group can be obtained (with some adjustment for population growth in the preceding year) from the regular census enumeration. Census officials, however, immediately discerned that the mortality data were underreported by approximately 40 percent. Life tables could only be constructed by making large (and ultimately unknowable) adjustments to the number of deaths reported at each age (see, e.g., Elliott 1874's "approximate" life table for the 1870 population). Differential mortality could be examined only by assuming no differentials in undercounts. Census officials clearly believed, however, that the undercount varied by region, urban/rural residence, and between long and recently settled states. J. D. B. De Bow (1855, 8), superintendent of the 1850 census, contended that state differentials in death rates "show not so much in favor of or against the health of either, as they do, in all probability, a more or less perfect report of the marshals. Thus it is impossible to believe Mississippi a healthier State than Rhode Island." Despite this disappointment, and the urging of some census officials to drop the expensive undertaking, the mortality information was deemed useful enough to continue its collection and publication through the 1900 census. More questions were added and, beginning in 1880, the information was supplemented with death records from states with available registration data (Condran and Crimmins 1979).

Retrospective mortality data were undercounted for several reasons. Most obviously, solitary households left no one behind to report the death to an enumerator. The death of a household member of a larger family, especially the household head, often led to the dissolution of the household.

Respondent error also led to undercounting. Deaths of infants and the elderly were underreported, and deaths occurring 6–12 months prior to the census enumeration were less likely to be reported than deaths occurring 0–6 months prior to the count (Condran and Crimmins 1979; Ferrie 2003). In an early comparison of death reporting between the 1880 census of mortality and the early death registration states of Massachusetts and New Jersey, J. S. Billings (1885, xlii) observed that "the proportion of deaths omitted in the enumerators' returns increases in a tolerably regular manner as we go back in time from the date of enumeration." Billings calculated that census undercounting of deaths in the 1880 census increased from about 17 percent of all registered deaths 0–6 months prior to the census to 30 percent of deaths registered 6–12 months prior.

Despite severe underenumeration, researchers have made creative use of the mortality censuses. By matching deaths registered in the DRA to deaths registered by the mortality censuses, Condran and Crimmins (1980, 188–90) were able to estimate undercounts in both sources and make a more accurate comparison of urban and rural mortality. Ferrie (2003) used surviving original manuscript returns from the 1850 and 1860 mortality census to link decedents to their household of origin and was thus able to investigate mortality differentials by age, occupation, wealth, nativity, migration status, and household size. The use of linked microdata allowed Ferrie (1996) to make another important innovation: by relying only on deaths reported in the six months prior to the census, he was able to significantly reduce respondent recall error and construct adult life expectancy estimates for white males by region, urban/rural residence, and nativity. The results suggest a substantial advantage in life expectancy at age 20 for white males living in rural areas and for native-born males.

Haines (1998) has made the most significant attempt to use the mortality censuses to construct life tables. He began by observing that the underreporting of deaths for individuals age 5–9, 10–14, and 15–19 appeared to be small. By fitting age-specific mortality rates for these age groups to model life tables, Haines was able to avoid relying on age groups experiencing substantial underreporting of deaths and to construct life tables for the white, black, and total populations by sex for each census year between 1850 and 1900. These tables are clearly superior to their predecessors and a major step forward in our understanding of late nineteenth-century mortality. Despite some concern about regional and temporal differences in undercounting, mortality data were collected for the entire nation. Thus, with the exception of Meech's (1898) 1830–60 life table, Haines's tables can be considered the only nationally representative life tables for the nineteenth-century United States. The availability of life tables every 10 years between 1850 and 1900 also filled many of the gaps between existing life tables. Contrary to most prior assumptions, Haines's life tables indicated that mortality did not begin its secular decline until relatively late in the century. Life expectancy at birth was variable without trend between 1850 and 1880—ranging between 38.3 and 44.0 years for both sexes combined. Between 1880 and 1900, however, life expectancy at birth increased from 39.4 to 47.8 years (U.S. model, both sexes combined).

Researchers relying on the Haines (1998) life tables need to be aware of a few potential problems with their interpretation and use. First, as Haines noted, the life tables represent mortality conditions only in the year preceding each decennial census and thus may not be representative of the period or decade in which they nominally represent. Haines's 1850 life table, for example, like Jacobson's 1850 life table, may overstate mortality because of the 1849 cholera epidemic. Interpolating between Haines's life tables for the intercensal periods between 1850 and 1880 suggests that individuals living in the 1860s enjoyed the period's lowest mortality. The opposite is likely true. During the 1860s the United States suffered four years of the Civil War, a major and prolonged depression in the postwar South, and, in 1867, another major epidemic of cholera. The war alone is believed to have resulted in the death of approximately 8 percent of white men aged 13 to 43 in 1860 (Vinovskis 1989). Finally, users of the Haines life tables should also be aware that the shape of age-specific mortality rates are strongly influenced by the Haines's choice of models: model "west" of the Princeton regional model life tables and a "U. S model" derived from the 1900–1902 DRA life table. As discussed below, there is evidence that these models fail to accurately describe the age profile of mortality in the nineteenth-century United States, particularly for women in their childbearing years. Despite these qualifications, Haines's life tables are a major point of reference for the latter half of the nineteenth century.[4]

The only studies of life expectancy prior to 1850 approaching the geographic coverage of the Haines life tables are genealogical-based estimates of adult life expectancy by Kunze (1979), Fogel (1986), and Pope (1992), and mean age at death estimates by Kasakoff and Adams (1995). Because genealogies observe individuals from birth to death, cohort life expectancies are easily calculated. Period estimates can also be made by observing deaths and years of exposure over a given interval, typically a decade. Decennial life expectancy estimates thus reflect mortality over the entire decade, not just a single year. Because individuals are followed over time and space, genealogical data allow the application of event-history methods and more sophisticated analyses. Kasakoff and Adams (2000), for example, were able to examine the impact of migration on subsequent mortality. There are several drawbacks to the use of genealogical data for estimating mortality, however, including substantial underreporting of infant and childhood deaths (thus limiting estimates to adult life expectancy), underreporting of female deaths, a bias toward larger and longer-lived families, lack of coverage of the nation's black and foreign-born populations, small sample sizes for early birth cohorts, a bias toward married individuals who reproduce, and a bias toward families originating in the Northeast and living in the North. Kasakoff

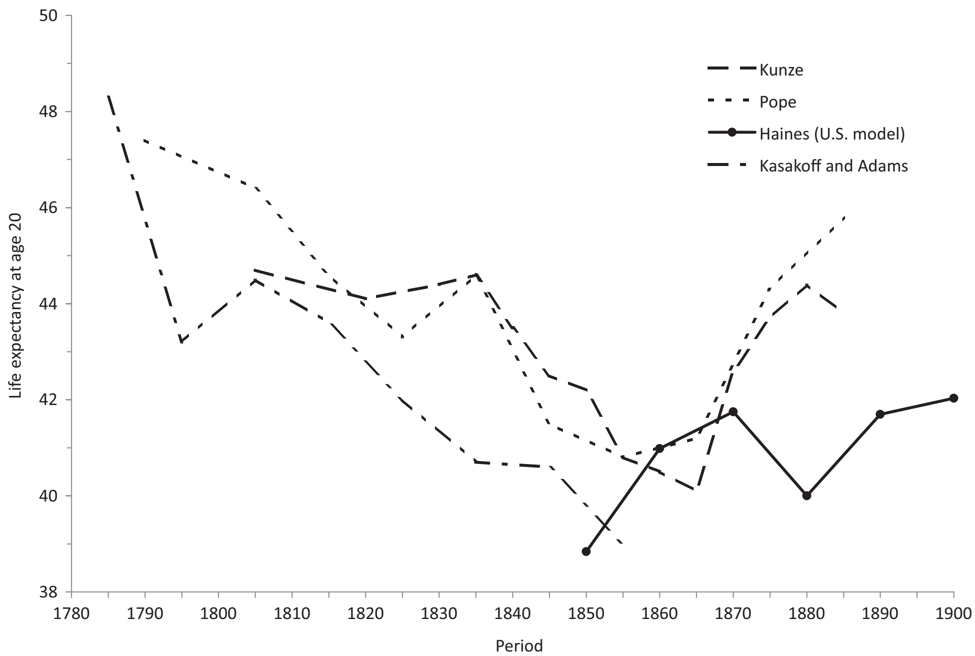

FIGURE 1. Male life expectancy at age 20.

and Adams's data set, for example, was drawn from nine published genealogies of families whose ancestors settled in seventeenth-century New England. Although nineteenth-century descendents of the nine families can be found in all parts the nation, they were primarily located in the nation's northern census regions. Kunze's (1979) and Pope's (1992) data sets were drawn to be more representative of the regional distribution of the United States population. Although not perfectly representative, the geographic coverage of both samples is reasonably representative of the overall population.

Figure 1 plots estimates of white male life expectancy at age 20 by Kunze (1979), Pope (1992), and Haines (1998), and mean age at death estimates for white males known to survive to age 20 by Kasakoff and Adams (1995).[5] Four observations can be made. First, the three genealogical studies report very high adult male life expectancies in the late eighteenth and early nineteenth centuries; if the estimates are correct, adult life expectancies in the United States at the turn of the nineteenth century were the highest in the world and were not again exceeded in the United States until circa 1920, approximately four decades after the onset of secular mortality decline. Second, life expectancy estimates by Haines (1998) are about three years lower, on average, than those reported in the genealogical studies in the decades in which they overlap and can be reliably compared. Third, although there is much variation in each study's sources, methods, and results, it is nonetheless clear from figure 1 that the genealogical-based studies support Haines's contention that mortality did not begin its secular decline until late in the century. Finally, all three genealogy-based studies suggest a significant increase in mortality in the antebellum era, especially in the three decades between 1830 and 1860. White male life expectancy at age 20 was approximately six years lower at mid-century than it was in the late eighteenth century.

If correct, a substantial mid-century increase in mortality represents a paradox; based on an assessment of the expected impact of urbanization, public health, and economic growth, Easterlin (1977) had hypothesized a substantial mortality decline before 1880. Although urbanization increased during the period, facilitating the spread of infectious disease and higher mortality, Easterlin noted that the percentage of the United States population living in urban areas remained modest until late in the century. The urban population, for example, was just 28.2 percent in 1880.[6] Given an expected 10-year urban-rural differential in life expectancy—an approximate differential suggested by several studies—and assuming a negligible role of public health before 1880, Easterlin estimated that urbanization between 1800 and 1880 reduced life expectancy at birth 2.1 years, all else being equal. The negative effect of urbanization, however, was more than compensated for by increases in the standard of living. Real national income per capita increased dramatically in the period before 1880, leading to significant improvements in diet and housing.[7] By assuming a theoretical relationship between life expectancy and per capita income suggested by cross-sectional national data for the twentieth century (Preston

1975), Easterlin estimated that growth in real income in the period 1800–80 should have increased life expectancy by 14 years. Together with the negative impact of urbanization, Easterlin's model suggested that life expectancy at birth increased 11.9 years between 1800 and 1880.

Although a reasonable theoretical argument for declining mortality, Easterlin conceded serious doubts in estimates of national income in the period before 1840, the appropriateness of using the relationship between income and life expectancy in the twentieth century to infer the relationship a century earlier, the possibility that public health worsened between 1800 and 1880, and the need for more empirical research. Given these doubts, new estimates documenting a mid-nineteenth-century mortality increase cannot be dismissed on theoretical grounds. Moreover, indirect support for an "antebellum paradox" of increasing mortality during a period of strong economic growth is provided by new research on the anthropometric history of the nineteenth-century United States. Fogel (1986, 464–67) first called attention to the positive long-run correlation between cohort life expectancy at age 10 and the final achieved heights of white men. Both series decline in the early to mid–nineteenth century and increase late in the century. Accumulating evidence from other sources confirms a substantial decline in male height for cohorts born in the mid–nineteenth century. Dora L. Costa and Richard H. Steckel (1997, 72), for example, documented a decline in stature among native-born white males from a mean of 173.5 centimeters in the 1830 birth cohort to 169.1 in the 1890 cohort, followed by a substantial and sustained increase in heights for cohorts born in the twentieth century. While identification of the causes of the decline has been difficult—hypotheses include deteriorating diets, a worsening disease environment, the negative impact of early industrialization and urbanization, increasing rates of internal migration, and rising inequality—all researchers have agreed that heights declined significantly. In a recent investigation of the link between antebellum mortality, heights, and net nutrition, Michael R. Haines, Lee A. Craig, and Thomas Weiss (2003) have pointed to the importance of an increasing nationalization and internationalization of the disease environment. Regardless of the ultimate causes, the positive correlation between stature and life expectancy is additional evidence that the decline of life expectancy in the mid-nineteenth century reported by recent studies reflects a real increase in mortality.

There are ample reasons to remain skeptical of the overall level of life expectancy reported by the genealogical studies and the size of the suggested decline, however. Genealogical records suffer from two types of bias: a selection bias incurred by selecting data from demographically successful, native-born families, and a censoring bias incurred by excluding individuals without complete birth and death information from the analysis. Although these biases act in opposite directions—selection bias causes life expectancy estimates to be biased upwards while the censoring bias typically imparts a downwards bias—it is unlikely that they counteract each other perfectly and consistently.[8]

Adult life expectancy estimates based on genealogical sources tend to be much higher than estimates based on other types of sources, suggesting that selection bias dominates. Between 1785 and 1814, graduates of Yale College—an elite New England population with nearly complete, high-quality demographic data—had a life expectancy at age 20 of 40.4 years; Kunze's (1979) and Pope's (1992) genealogical estimates for the same period are much higher, in the mid-to-upper 40s (Hacker 1996, 121). Adult life expectancies of other elite colonial populations were even lower than that enjoyed by Yale graduates and were especially low in the colonial South. Life expectancy at age 20 was 36.2 years for men graduating from Princeton College between 1709 and 1819; 34.7 years for Maryland legislators born between 1750 and 1764; and 31.7 years for South Carolina legislators born 1750–64 (Levy 1996; Hacker 1996). Even if we assume no significant socioeconomic status differentials in adult mortality, these studies suggest that genealogical sources overestimate male life expectancy at age 20 at the turn of the nineteenth century by 5–10 years or more. Daniel S. Levy (1996) indicates that lower life expectancy in the colonial South was rapidly disappearing by the late eighteenth century, however, suggesting that the overstatement of male life expectancy by genealogical sources was on the lower side of that range, perhaps six years in the last decade of the century.

The tendency of genealogical estimates to overstate adult male life expectancy appears to have been lower in the mid- and late nineteenth century. In the two periods where they can be compared—1850–60 and 1870–90—Kunze's (1979) and Popes' (1992) combined estimates of male life expectancy at age 20 are 2.73 years higher, on average, than Haines's (1998) estimates.[9] Male life expectancy estimates derived with two-census methods suggest a similar differential. Table 3 shows the results of applying the Samuel Preston and Neil Bennett's (1983) two-census method to the native-born white population enumerated in the 1850 and 1860 IPUMS censuses. The method assumes the population is closed to migration, a reasonable assumption for the native-born population of the nineteenth-century United States. Although the results may be biased by differential undercounting and the accuracy of age reporting in the two censuses, the resulting life table suggests that genealogical estimates overstate male life expectancy at age 20 in the 1850s by about 3.5 years. Unfortunately, substantial underenumeration of the 1870 census (see Anderson 1988, 78–82; Steckel 1991) limits comparison to the decade 1850–60.

The lower tendency of genealogical sources to overstate life expectancy in the mid- and late nineteenth century may be the result of greater migration censoring in the genealogical data. The opening of the trans-Appalachian west with the Treaty of Paris in 1783, the defeat of the Pan-Indian alliance in 1793, land reforms in the early nineteenth century, and the transportation revolution of the 1830s likely increased

TABLE 3. Application of the Preston-Bennett Census-Based Method to the Native-Born White Population by Sex: 1850–60

Start of age interval (x)	Population on June 1, 1850	Population on June 1, 1860	Aveage population	Intercensal growth rate ${}_5r_x$	Cumulated growth rate	Stationary population interval ${}_5L_x$	Stationary population above age x, T_x	Number surviving to age x in stationary population	Estimated life expectancy at age x,e_x
Males									
0	1, 423, 462	2, 053, 500	1, 738, 481	0.0366	—	—	—	—	—
5	1, 316, 436	1, 698, 039	1, 507, 238	0.0254	0.06358	1, 606, 187	—	—	—
10	1, 147, 038	1, 446, 005	1, 296, 522	0.0231	0.18503	1, 560, 040	14, 080, 177	316, 623	44.5
15	956, 661	1, 233, 984	1, 095, 323	0.0254	0.30647	1, 488, 128	12, 520, 137	304, 817	41.1
20	830, 860	1, 055, 632	943, 246	0.0239	0.42987	1, 449, 818	11, 032, 008	293, 795	37.6
25	654, 370	855, 794	755, 082	0.0268	0.55671	1, 317, 558	9, 582, 191	276, 738	34.6
30	548, 139	678, 327	613, 233	0.0213	0.67697	1, 206, 791	8, 264, 632	252, 435	32.7
35	452, 270	584, 639	518, 455	0.0257	0.79433	1, 147, 321	7, 057, 842	235, 411	30.0
40	372, 137	471, 681	421, 909	0.0237	0.91767	1, 056, 228	5, 910, 521	220, 355	26.8
45	310, 999	400, 900	355, 950	0.0254	1.04031	1, 007, 371	4, 854, 293	206, 360	23.5
50	256, 448	332, 500	294, 474	0.0260	1.16861	947, 478	3, 846, 923	195, 485	19.7
55	165, 102	225, 940	195, 521	0.0313	1.31185	725, 977	2, 899, 444	167, 346	17.3
60	146, 113	194, 447	170, 280	0.0286	1.46160	734, 392	2, 173, 467	146, 037	14.9
65	93, 573	121, 785	107, 679	0.0263	1.59881	532, 703	1, 439, 075	126, 710	11.4
70	61, 019	77, 378	69, 199	0.0237	1.72397	387, 976	906, 372	92, 068	9.8
75	35, 364	46, 194	40, 779	0.0267	1.85003	259, 355	518, 395	64, 733	8.0
80	20, 515	24, 696	22, 606	0.0185	1.96310	160, 982	259, 040	42, 034	6.2
85+	10, 913	13, 798	12, 356	0.0234	2.07147	98, 059	98, 059	—	—
Females									
0	1, 383, 318	2, 021, 279	1, 702, 299	0.0379	—	—	—	—	—
5	1, 266, 758	1, 674, 058	1, 470, 408	0.0279	0.06964	1, 576, 457	—	—	—
10	1, 106, 856	1, 377, 428	1, 242, 142	0.0219	0.19391	1, 507, 943	13, 610, 916	308, 440	44.1
15	968, 287	1, 260, 281	1, 114, 284	0.0263	0.31437	1, 525, 903	12, 102, 973	303, 385	39.9
20	812, 808	1, 070, 750	941, 779	0.0275	0.44906	1, 475, 614	10, 577, 071	300, 152	35.2
25	634, 318	791, 899	713, 109	0.0222	0.57333	1, 265, 175	9, 101, 457	274, 079	33.2
30	505, 521	638, 730	572, 126	0.0234	0.68718	1, 137, 446	7, 836, 282	240, 262	32.6
35	418, 015	531, 753	474, 884	0.0240	0.80572	1, 062, 940	6, 698, 835	220, 039	30.4
40	349, 205	447, 572	398, 389	0.0248	0.92783	1, 007, 534	5, 635, 895	207, 047	27.2
45	288, 765	355, 900	322, 333	0.0209	1.04204	913, 814	4, 628, 361	192, 135	24.1
50	238, 868	309, 831	274, 350	0.0260	1.15924	874, 487	3, 714, 547	178, 830	20.8
55	161, 663	213, 946	187, 805	0.0280	1.29421	685, 129	2, 840, 060	155, 962	18.2
60	137, 924	190, 055	163, 990	0.0320	1.44429	695, 124	2, 154, 931	138, 025	15.6
65	91, 750	116, 072	103, 911	0.0235	1.58311	506, 056	1, 459, 807	120, 118	12.2
70	63, 448	82, 476	72, 962	0.0262	1.70737	402, 343	953, 752	90, 840	10.5
75	36, 275	48, 482	42, 379	0.0290	1.84534	268, 267	551, 409	67, 061	8.2
80	20, 104	24, 198	22, 151	0.0185	1.96410	157, 903	283, 142	42, 617	6.6
85+	15, 154	16, 395	15, 775	0.0079	2.07183	125, 239	125, 239	—	—

*Source:*1850–60 IPUMS samples from Ruggles before the et al. Steven Ruggles, Matthew Sobek, J. Trent Alexander, Catherine A. Fitch, Ronald Goeken, Patricia Kelly et al., *Integrated Public Use Microdata Series: Version 4.0* (Minneapolis, MN: Minnesota Population Center, 2009) [producer and distributor].

the level and typical distance of internal migration. In the seven decades between 1790 and 1860 the area of the United States increased from 891,364 to 3,021,295 square miles and the number of states from 16 to 33, with the greatest increases between 1840 and 1860 (Anderson 1988, 241, 246). The mean center of population moved further west in the two decades between 1840 and 1860—135.4 miles—than in any other comparable period in United States history (U.S. Bureau of the Census 1921, 34). Although we cannot be sure of the size and timing of the effect—Kunze (1979) and Pope (1992) do not report the percentage of their study populations with missing death dates by decade—migrants are more likely to be lost from observation. Without adequate attempts to adjust the population at risk, an increase in the percentage of right-censored cases would bias life expectancy estimates downwards, all else being equal.[10]

TABLE 4. Suggested Best Estimates for Male Life Expectancy at Age 20 (e_{20})

Period	Male e_{20} from genealogical-based studies	Suggested correction factor	Adjusted male e_{20}
1790–99	47.4	6.0	41.4
1800–09	45.8	–5.5	40.3
1810–19	44.6	–4.9	39.7
1820–29	44.1	–4.4	39.7
1830–39	43.8	–3.8	39.9
1840–49	42.6	–3.3	39.3
1850–59	41.2	–2.7	38.4
1860–69	40.8	–2.7	38.0
1870–79	43.7	–2.7	41.0
1880–89	45.0	–2.7	42.2
1890–99	n.a.	n.a.	43.2 [a]

[a]Interpolated from the 1880–89 adjusted estimate and a weighted average of the 1900–1902 DRA and rural DRA life tables.
Sources: Kent Kunze, *The Effects of Age Composition and Changes in Vital Rates on Nineteenth-Century Population Estimates from New Data* (Salt Lake City, UT: Department of Economics, University of Utah, 1979); and Clayne L. Pope, "Adult Mortality in America before 1900: A View from Family Histories," in *Strategic Factors in Nineteenth Century American Economic History: A Volume to Honor Robert W. Fogel*, ed. Claudia Goldin and Hugh Rockoff, 267–96 (Chicago: University of Chicago Press).

Selection bias may also have been less important in the nineteenth century than in the eighteenth century. If selection bias is a function of the propensity of a long-lived ancestor to produce a large number of descendents—thus increasing the odds of producing a future genealogist—the life expectancy of earlier birth cohorts is more critical to the subsequent number of descendants than that of later, larger cohorts, where we can expect more heterogeneity. Put another way, the chances that a couple will produce any descendents beyond a few generations is low if their mortality or the mortality of their children and grandchildren is high. If mortality is low in the first few generations, however, the chances are very high that there will be thousands of descendants (and many potential genealogists) regardless of the level of mortality in subsequent generations (for a general discussion of these issues with regards to Chinese demographic history see Zhao 2001).

Despite concerns about selection and censoring biases, it is clear from recent studies that mortality increased significantly after 1830 and remained relatively high until the 1870s, at which point it began its long and sustained decline. Although genealogical-based estimates of male life expectancy are biased upwards, especially in the eighteenth and early nineteenth centuries, they represent our best source for decennial trends in life expectancy between 1790 and 1890. With care, the estimates can be combined and adjusted to construct a reasonable series of adult life expectancies.

Table 4 attempts such a series by averaging the Kunze (1979) and Pope (1992) estimates of male life expectancy in each decade and adjusting the combined estimates by a correction factor suggested by comparisons with other studies. The second column shows the average Kunze and Pope estimate of male life expectancy at age 20 for each decade between 1790 and 1890.[11] The third column shows a suggested correction factor for each of these decades: –6 years in the late eighteenth century (suggested by comparisons with the graduates of Yale College and other special populations) and –2.73 years in the mid- to late nineteenth century (suggested by comparison to Haines's [1998] life tables). The correction factor is interpolated between the 1790s and the 1850s, corresponding to suspected trends in regional migration. The adjusted male life expectancy estimates are shown in the fourth column. Because Pope's and Kunze's genealogical estimates for adult life expectancy end with the 1880–89 decade, the suggested male estimate for the period 1890–99 was obtained by interpolating between the 1880–89 estimate and an estimate obtained from the 1900–1902 overall and rural DRA life tables, weighted to reflect the national level of urbanization. (The 1900–1902 DRA life tables and their weighting to reflect national levels of urbanization is subsequently described in more detail.)

Correction factors for the early part of the century are clearly larger and more speculative than those in the second half of the nineteenth century. Indirect evidence suggests that they are approximately correct, however. Given the age structure of the population reported in the United States census of 1800, the adjusted estimates in the fourth column imply a crude birth rate for the white population of 51.5 births per 1,000 inhabitants. The unadjusted estimate, however, would imply a crude birth rate of 45.6 per 1,000, whereas a 2.73-year adjustment would imply a birth rate of 47.7 per 1,000. Contemporary observers and twentieth-century demographers have agreed that the birth rate at the turn of the nineteenth century was between 50 and 57 per 1,000, strongly suggesting that the six-year adjustment is justified (Grabil, Kiser, and Whelpton 1958, 5; McClelland and Zeckhauser 1982, 71).[12] Although based in part on trends in internal migration and the known impact of migration censoring on mortality estimates, and in part on the observed bias in the genealogical-based estimates of life expectancy compared to other sources, the linear interpolation of the adjustment factor between the 1790s and 1850s is also speculative. As a result, life tables constructed from these estimates will have a larger margin of error than life tables constructed from estimates for the latter part of the century.

The adjusted estimates shown in the fourth column suggest that male life expectancy at age 20 declined approximately three years between 1790–99 and 1850–59. Male life expectancy continued to decline in the 1860s, due largely to the impact of the Civil War. Thereafter, life expectancy began its long-term, sustained increase. It is unlikely that mortality was under significant human control until circa 1880, however. The adjusted series suggests that male life expectancy at age 20 did not exceed its level in the late eighteenth century until the 1880s.

The suggested series indicates a more moderate decline in antebellum life expectancy than the six-year decline suggested by the unadjusted genealogical estimates. The decline is still large, however, and remains a puzzling aspect of nineteenth-century United States demographic history. The suggested revisions shown here do not negate scholars' characterization of the decline as an "antebellum paradox" or the need for more research on the causes of declining health and longevity during a period of rapid economic growth.

Sex Differentials in Adult Life Expectancy

Estimating female life expectancy at age 20 using genealogical records is a major challenge. Because women appear less often in public records and change their surname at marriage, they disappear from observation more frequently than men. And because genealogical records do not record when right-censored individuals exit observation, female estimates of life expectancy are based on fewer cases and subject to more censoring biases than male estimates.[13]

Difficulties determining when women entered and exited observation and small sample sizes in each decade likely explain the highly variable sex differentials in adult life expectancy reported by Kunze (1979) and Pope (1992) (see table 1). Pope reported that women experienced a 1.6-year advantage in life expectancy at age 20 in the 1820s and a 4.4-year disadvantage in the 1840s. Kunze reported that females had a 3.4-year advantage in the period 1830–34 and a 2.3-year disadvantage in 1835–39. Such rapid shifts in sex differentials in life expectancy are likely spurious and related to poor data quality.

Unfortunately, there are few studies of eighteenth- and early nineteenth-century female life expectancy that can be used to evaluate potential biases in Pope's (1992) and Kunze's (1979) estimates. Female life expectancy estimates derived using other sources and methods (e.g., estimates from community-based reconstruction studies) are also based on incomplete data and subject to substantial selection and censoring biases (see Hacker 1997, for a summary of existing studies and discussion of potential biases). Life expectancy of women married to Yale graduates at age 20, for example, was five years lower in the late eighteenth than the estimates reported by Kunze (1979) and Pope (1992). Although the difference is approximately equal to the difference observed between the genealogical estimates and the life expectancy of Yale graduates, more than one-in-four Yale wives had an unknown date of death, rendering an assessment of bias in the genealogical estimates uncertain. Given different assumptions about the mortality experiences of women with a missing death date, the life expectancy of Yale wives at age 20 may have been one year higher or lower (Hacker 1996, 83, 98). Much higher proportions of missing data and margins of error characterize other late eighteenth- and early nineteenth-century estimates of female life expectancy.

For the late nineteenth century, Kunze's (1979) and Pope's (1992) estimates of the female life expectancy can be compared with Haines's (1998) estimates. The comparison indicates that Kunze's (1979) and Pope's (1992) combined estimates for white females at age 20 are slightly lower (–0.36 years) than Haines's estimates in the years in which they can be reliably compared. This is in sharp contrast to the comparison with Haines's estimate for white males, in which the genealogical-based estimates were substantially higher (2.73 years at age 20). Given the high proportion of missing death records for women in genealogies, the difficulties determining when women entered and existed the at-risk population, and the highly variable sex differentials in life expectancy reported by Kunze and Pope, it is tempting to conclude that this discrepancy is due entirely to bias in estimating female life expectancy from genealogical data. Some portion of the difference in the male and female comparisons with Haines's life table estimates may be caused by Haines's choice of a model life table system, however. This possibility is explored in the subsequent section examining age patterns in nineteenth-century mortality.

Regardless of the ultimate cause, poor data quality, inconsistent results, and the lack of an independent assessment of potential bias strongly suggests that determination of the level of and trend in female life expectancy is best inferred from male estimates. This section discusses sex differentials in nineteenth-century life expectancy, suggests a best estimate for the differential at age 20 in each decade and calculates the resulting series of female life expectancy from the adjusted male estimates shown in table 4. The sex differential is assumed to be constant before 1860, after which fertility and mortality decline are assumed to have contributed to more rapid female gains in life expectancy relative to male gains (see Preston 1976, chap. 6, for a discussion of the impact of mortality decline on sex differentials in mortality). Estimates are made separately for the 1860s to account for excess male mortality during the Civil War.

The best estimate of the sex differential in life expectancy for the period before 1860 and the best estimate for each decade after 1870 are not obvious from existing studies of nineteenth-century U.S. mortality. Kunze's (1979) and Pope's (1992) estimates suggest a male advantage in life expectancy at age 20 while Haines's life tables suggest a female advantage. On average, the combined Pope and Kunze estimates of male and female life expectancies at age 20 indicate a 0.9-year male advantage before 1860. For census years 1850 and 1860, Haines's U.S. model life tables suggest an average female advantage in life expectancy at age 20 of 1.1 years (Haines's life tables based on Princeton model west life tables indicate a 2.9-year female advantage).

These contrasting results persist in the postwar era. Kunze's (1979) and Pope's (1992) results indicate that males enjoyed a 2.4-year advantage, on average, in the 1870s and 1880s while Haines's U.S. model life tables indicate a 1.3-year female advantage (2.7 years using model West). At the beginning of the twentieth century, the 1900–1902 DRA life table shows a 1.6-year female advantage in life expectancy at age 20, which is in close agreement with Haines's U.S. model. The close agreement is not surprising, of course; Haines's U.S. model life tables are based on the age pattern of mortality in the 1900–1902 DRA. The life table constructed for the rural parts of the 1900–1902 DRA, however, shows a female advantage in life expectancy at age 20 of just 0.1 years, closer to the implied sex differential in the combined Pope and Kunze estimates.

The different sex differentials in adult life expectancy observed in the overall and rural DRA life tables hint that males may have enjoyed higher adult life expectancies in the more rural past. Such a conclusion is supported by the demographic literature on nineteenth-century European populations.[14] A recent comparative study of mortality in rural villages in eighteenth- and nineteenth-century Europe and Asia (the Eurasia Population and Family History Project), for example, reports lower female lower life expectancy at age 25 in three of the four European study areas. Sex differentials in life expectancy at age 25 was –2.3-years for Sart, Belgium (a 2.3-year female disadvantage relative to males); –1.0 years for Casaluidi, Italy; –2.8 years for Madregolo, Italy; and 0.7 years for Scanian parishes in Sweden, for an unweighted average of –1.4 years (Campbell, Lee, and Bengtsson 2004, 66). Lower female life expectancy at age 25 resulted from a remarkably consistent pattern of higher female mortality during prime childbearing ages across study populations, suggesting that maternal mortality and maternal depletion played a large role in the consistent pattern (Alter, Manfredini, and Nystedt 2004). The pattern is characteristic of mortality in national populations with life expectancy below 45 and suggestive of higher female mortality from pulmonary tuberculosis, other infectious diseases, and maternal causes (Preston 1976, 91).

Some evidence suggests that females in rural areas of nineteenth-century Europe suffered higher rates of infectious disease relative to males than females in urban areas. Dominique Tabutin and Michel Willems, for example, cite evidence that excess female mortality and susceptibility to respiratory diseases such as tuberculosis were more pronounced in rural areas (cited in Alter, Manfredini, and Nystedt 2004). Excess female mortality extended over a greater range of ages and was much higher in England's 63 "healthy districts"—mostly rural districts with crude death rates below 17 per 1,000—than in the 1838–54 English Life Table (Woods 2000, 187). According to Shelia Ryan Johansson, a probable reason for the higher incidence of tuberculosis among females and higher rates of female mortality in rural areas of Victorian England was lower nutritional status. Agricultural societies in the past, she observed, routinely discriminated against females by reserving most food and the vast majority of meat for husbands and sons. Industrialization and the ability of women to participate in the paid labor force eventually ended this nutritional discrimination (Johansson 1977). Higher fertility is another possible reason for higher female mortality in rural areas. Although maternal mortality rates were low relative to mortality rates from tuberculosis—most nineteenth-century estimates suggest that maternal mortality averaged between 5 and 10 maternal deaths per 1,000 live births (Kippen 2005)—higher rates of nuptiality and martial fertility in rural areas increased the cumulative risk of maternal mortality. Perhaps more importantly, pregnancy and lactation imposed greater nutritional demands on women and reduced cell-mediated immunities, increasing the risk of contracting tuberculosis and other opportunistic infections.[15]

Unfortunately, with the exception of Kunze's (1979) and Pope's (1992) studies, estimates of sex differentials in life expectancy for the nineteenth-century United States are based on highly urban, low-fertility populations such as Massachusetts in the late nineteenth century, the 1900–1902 DRA, or, like Haines's (1998) life tables, are derived from models based on these populations. The 1850–60 Preston-Bennett life table (table 3), however, avoids this urban, low-fertility bias by relying on the national native-born white female population in the 1850 and 1860 IPUMS samples. The results suggest sex differentials in life expectancy similar to Kunze's (1979) and Pope's (1992) genealogical-based estimates. At age 15, the sex differential in life expectancy was –1.2 years, rising to –2.3 years at age 20. The male advantage in life expectancy lasted until age 35. Thereafter, females enjoyed a slight advantage in expected remaining years of life.

Together, the results from eighteenth- and nineteenth-century European populations and the results indicated by the 1850–60 Preston-Bennett life tables for native-born whites suggest that the overall average 0.9-year male advantage in life expectancy at age 20 reported by Kunze and Pope for the period 1780–1859 was approximately correct.[16] As indicated by the 1900–1902 DRA life tables, however, a female

TABLE 5. Suggested Best Estimates for Female Life Expectancy at Age 20

Period	Adjusted male e_{20}	Suggested sex differential (female-male)	Suggested female e_{20}
1790–99	41.4	–0.9	40.5
1800–09	40.3	–0.9	39.4
1810–19	39.7	–0.9	38.8
1820–29	39.7	–0.9	38.8
1830–39	39.9	–0.9	39.0
1840–49	39.3	–0.9	38.4
1850–59	38.4	–0.9	37.5
1860–69	n.a.	n.a.	38.9 [a]
1870–79	41.0	–0.6	40.4
1880–89	42.2	0.0	42.2
1890–99	43.2	0.6	43.8

[a]Average of period estimates from 1850–59 and 1870–79. See text.
Sources: Kent Kunze, *The Effects of Age Composition and Changes in Vital Rates on Nineteenth-Century Population Estimates from New Data* (Salt Lake City, UT: Department of Economics, University of Utah, 1979); and Clayne L. Pope, "Adult Mortality in America before 1900: A View from Family Histories," in *Strategic Factors in Nineteenth Century American Economic History: A Volume to Honor Robert W. Fogel*, ed. Claudia Goldin and Hugh Rockoff, 267–96 (Chicago: University of Chicago Press).

advantage in life expectancy at age 20 had emerged by the turn of the twentieth century. If the overall and rural 1900–1902 life tables are weighted and combined to approximate the urban percentage of the national population, the female advantage in life expectancy at age 20 was 0.9 years in 1900.[17]

Table 5 suggests best estimates of female life expectancy at age 20 between 1780 and 1860 by assuming a fixed 0.9-year advantage in male life expectancy. As shown in the third column, the sex mortality differential was assumed to shift in favor of females in a linear fashion between the 0.9 female disadvantage in life expectancy in the period before 1870 and the 0.9-year advantage in female life expectancy suggested by the weighted 1900–1902 DRA life tables. Although somewhat speculative, the linear shift from a male advantage to a female advantage in life expectancy between 1870 and 1900–1902 is consistent with known changes in sex mortality differentials accompanying mortality decline, the epidemiological transition, and fertility decline. The decline in pulmonary tuberculosis, in particular, likely led to more rapid declines in female mortality relative to male mortality (Preston 1976, chap. 6). Because excess male mortality during the Civil War likely affected sex differentials in mortality, the female estimate of life expectancy in the period 1860–69 was obtained by averaging the adjusted female life expectancy in the 1850s and 1870s. Suggested best estimates of female life expectancy at age 20 are shown in the fourth column.

The Age Profile of Nineteenth-Century Mortality

Mortality varies with age in a consistent pattern, sometimes characterized as a "U" or "J" shape, across a wide range of mortality levels. Mortality rates are very high in infancy, drop rapidly in childhood, reach their lowest level in late childhood and adolescence, and then begin to increase in a fairly regular manner with age. Because of this consistency, demographers have long sought to model mortality as a function of age and overall mortality. Among other uses, an accurate model would make it possible to identify deviations in empirical data from model patterns (suggestive of particular conditions or poor data quality), to gain insight into environmental and behavioral factors that may determine deviations, and to construct life tables from poor data, partial data, or even a single parameter (Preston, Heuveline, and Guillot 2001, 191–2). With an accurate model, for example, it would be possible to generate decennial life tables from the estimates of adult life expectancy suggested in tables 4 and 5. Choice of model, however, involves some guesswork and is a potential source of substantial error.

Three basic approaches have been used to model the age pattern of mortality: mathematical approaches that represent mortality as a function of age, tabular approaches that show expected patterns of age-specific mortality rates and other life table parameters at different mortality levels, and a combination of the first two approaches that uses a mathematical function to relate mortality in a given population to a tabulated standard population (Preston, Heuveline, and Guillot 2001, 192–201). Early attempts to describe the relationship between mortality and age with a single mortality function were unsuccessful (see Woods 2000, 170–90 for a discussion of nineteenth-century attempts to specify the "laws of vitality"). For a variety of reasons, including changes in behaviors and in the leading causes of death (e.g., smoking and cancer), the age pattern of mortality varies enough across time and space that a simple mathematical model is not practical. An attempt by Larry Heligman and John H. Pollard (1980), for example, required a complex equation with eight parameters to model the age profile of mortality from infancy to old age.

The second approach to modeling age patterns of mortality has been the publication of model life table systems—sets of "model" life tables at different levels of morality. The most popular set of model life tables, the Princeton regional models, were published by Princeton demographers Ansley J. Coale and Paul Demeny in 1966 and revised in 1983 (Coale, Demeny, and Vaughan 1966). Coale and Demeny examined empirical data from 326 historical and contemporary populations. From the 192 life tables deemed reliable, Coale and Demeny identified four regional patterns, which they used to construct four "families" of model life tables.

In the 1983 revision, male and female life tables are shown at 25 different levels of mortality, ranging from level 1 (female life expectancy at birth equal to 20 years) to level 25 (female life expectancy at birth equal to 80 years) for each of the four regional patterns. Intermediate levels are easily obtained by interpolation. The four groups closely conform to four regions of Europe, which was the primary source of the life tables. The *north model* is based largely on life tables from Scandinavian countries. It is characterized by low infant mortality and low mortality at older ages. The *east model* is based on life tables from Eastern Europe and is characterized by high infant mortality. The *south model* is based mostly on tables from Southern Europe and is characterized by high mortality under age 5 and above age 65 and low mortality between age 40 and 60. The *west model* is more of a residual group and is based on the largest number of life tables, including tables from Western Europe, the United States, Canada, Australia, New Zealand, and Japan. Other model life table systems—including those created by the United Nations—have been created for developing countries in Asia, Africa, and Latin America, where different environments and causes of death lead to different patterns of mortality than are found in Coale and Demeny's European-dominated system (United Nations 1982).

For populations with poor vital registration data, the choice of a model life table—and thus the assumed age profile of mortality—typically requires some guesswork. Colin Newell (1988, 165) notes that the "general, but not always helpful, rule is to use a [model life table] system which is flexible enough to let real features and irregularities through, but which is sufficiently robust to be unaffected by errors in the data." Because the age profile of mortality is largely the result of environmental and behavioral factors—which determine the distribution of causes of death and the level of mortality—most analysts try to rely on a model life table system based on data from a nearby region with a similar environment, behaviors, and level of mortality. U.S. historical demographers tend to rely on Coale, Demeny, and Vaughn's (1983) west model, which is based in part on historical life tables for the United States (including the 1900–1902 DRA life table). Robert V. Wells (1992), for example, used the west model to infer life expectancy at birth in colonial America from adult and child survival estimates reported in various studies. Suspecting probable underenumeration of infant deaths in the 1900–1902 DRA, Condran and Crimmins fitted mortality rates for the one to four age group to model West life tables in order to estimate life expectancy in urban and rural areas of the DRA (1980, 191). Where it can be compared to empirical data, the west model appears to be a good fit for the total and white populations of the early twentieth-century United States (Haines 1979, 197; Preston and Haines 1991, 66). Douglas Ewbank (1987), however, found that the age mortality profile of early twentieth-century black population of the United States more closely matched the United Nation's far east model life table.

TABLE 6. Implied Coale and Demeny Life Table Parameters when $e_{20} = 40$ years

Variable	West	North	East	South
Male				
Level	11.71	10.78	9.32	8.83
e_0	43.8	41.2	38.0	38.1
$1000q_0$	160.6	155.6	250.1	192.6
l_{20}	71597	66756	61927	61801
Female				
Level	9.71	8.68	8.07	7.41
e_0	41.8	39.2	37.7	36.0
$1000q_0$	166.7	161.7	235.7	193.0
l_{20}	67829	63075	61168	58004

Sources: A. J. Coale, P. Demeny, and B. Vaughan, *Regional Model Life Tables and Stable Populations* (New York: Academic Press, 1983).

Depending on the application, the choice of model can be important. Preston and Haines (1991) found that choice of regional model had very little impact on indirect estimates of child mortality in the 1900 census (1991, 64–67). Estimating infant mortality and life expectancy at birth from life expectancy at age 20, however, is problematic. Table 6 shows implied estimates of male and female life expectancy at birth, infant mortality rates, and the proportion of the population surviving to age 20 when male and female life expectancy is 40 years using the four Princeton regional models. The implied life expectancy at birth for males ranges from a high of 43.8 years in model West to a low of 38.0 years in the east model, a difference of nearly six years. Implied male infant mortality rates vary from a low of 156 per 1,000 in the north model to a high of 250 in the east model. Using the west model, nearly 72 percent of the population survived to age 20. In the east model, the percentage was less than 62 percent. Similar differences are observable for the female population. These differences illustrate the large potential error that can be incurred by relying on the wrong model to infer a complete life table from a single parameter.

The third approach to modeling the age profile of mortality, developed by William Brass (1971), uses a mathematical function to transform a standard life table. It thus represents a combination of the mathematical and tabular approaches. Brass observed that logits of the l_xs from any two life tables are related to one another by a linear relationship, making it possible to describe a set of logits in an observed or target population using the logits from the standard table and appropriate intercept and slope values. Briefly, the logit transformation of the l_x column is based on the equation:

$$\text{logit}(1 - l_x) = Y_x = 0.5\,\text{Log}\,e(1 - l_x/l_x), \qquad (1)$$

in which $l_0 = 1.0$. The logits of an observed population, $Y_{\text{Obs}}(x)$, are related to the logits of a standard population, $Y_s(x)$, by the linear equation:

$$\text{logit}\,(\text{Obs.}\, l_x) = Y_{\text{Obs}}(x) = \alpha + \beta\, Y_s(x). \qquad (2)$$

To fit an observed life table to a standard table, logits of the observed l_xs are plotted against the logits of the standard life table. A straight line is then fitted to the points (typically with simple linear regression or weighted regression techniques), and the intercept and slope of the line, α and β, are calculated. Once α and β are calculated, fitted logits can be computed from the standard logits, and the anti-logits can be taken to produce a set of fitted l_xs, as shown in the equation below:

$$\text{Fitted } l_x = \frac{1}{1 + e^{2Y_{Fit}(x)}}. \qquad (3)$$

When the intercept (α) equals 0 and the slope (β) equals 1, the standard table will be reproduced. Values of the intercept parameter greater than 0.0 will shift the level of mortality above the standard table and values less than 0.0 will shift the level of mortality below the standard table. The slope parameter determines the "tilt" of the table. A slope value greater than 1.0 indicates that infant and child mortality is lower relative to adult mortality than in the standard table, and a slope less than 1.0 indicates that infant and child mortality is higher relative to adult mortality than in the standard. It thus becomes possible to construct a family of related life tables from a standard life table by varying the intercept and slope parameters, calculating the anti-logits of the resulting values, and constructing the resulting life tables.

Although Brass (1971) suggested two sets of logits to use as a standard—a general standard and an African standard—any life table can be used and logits calculated directly from the l_x column. Appropriate choice of a standard table can preserve variations in the age profile of mortality that cannot be obtained by varying the slope and intercept parameters of a standard table, such as the level of older age mortality relative to mid-age mortality or the level of infant mortality relative to childhood mortality.[18] To construct his U.S. model life tables, for example, Haines (1979) relied on the 1900–1902 DRA life table as a standard table. With the help of available historical life tables from Massachusetts and other United States life tables of reasonable quality, Haines first estimated the impact of urbanization and time on the slope of the age mortality profile. While more urban environments increased infant and childhood mortality relative to adult mortality, the trend in the late nineteenth century was toward relatively lower levels of infant and child mortality (ibid., 303). From this relationship Haines determined the likely slope parameter in each census year between 1850 and 1900, effectively reducing the two-parameter logit model to a one-parameter model. The final intercept parameter was determined by fitting the age-specific death rates of children age 5–19 in the mortality censuses (ibid.).

Comparison of Haines's U.S. model life tables with the life tables constructed using the west model as a standard indicates that the U.S. model typically yields higher infant mortality rates, lower adult mortality rates, and lower life expectancy estimates at birth. In 1880, for example, Haines's U.S. model suggests an infant mortality rate of 0.214 for white males and a life expectancy at birth of 40.4 years. The life table constructed using a west model suggests an infant mortality rate of 0.180 and a life expectancy at birth of 40.9 years.

Arguably, the 1900–1902 DRA life table is a more appropriate standard for the nineteenth-century United States than a generic standard or even model West.[19] As noted in table 2, however, the DRA population was much more urban than the overall population in 1900, had a higher proportion of the population foreign born, had a lower proportion engaged in agriculture and had much lower fertility. The contrast is even greater with the overall population in the early and mid-nineteenth-century United States, which was overwhelmingly rural and had very high fertility. Although variation of the slope parameter can pick up some of the suspected impact of urbanization and time on the suspected age profile of mortality in the nineteenth century, the increase in urbanization and immigration, the decline in fertility and the agricultural sector of the economy, and the onset of the public health movement and epidemiological transition in the later part of the nineteenth century likely affected the distribution of causes of death and the age profile of mortality in more complex ways. It is likely, for example, that declining tuberculosis in the late nineteenth century had a significant impact on the mortality of young adults relative to infants and older adults, especially among females. Condran and Cheney (1982, 105) report that the decline in mortality from pulmonary tuberculosis explained 26.8 percent of the decline in mortality in Philadelphia between 1870 and 1900 and was overwhelmingly important in the decline in death rates at ages 20–39.

In addition to mortality decline, rapid fertility decline in the late nineteenth century (Hacker n.d.) likely had an impact on the age-specific mortality rates of females. Although maternal mortality rates were lower than typically imagined in the qualitative literature (Schofield 1986), repeated exposure to death in childbirth in high-fertility populations increased female mortality relative to male mortality during childbearing age.[20] Pregnancy may have been a significant risk factor in contracting tuberculosis, the leading killer of nineteenth-century Americans, and other opportunistic infections. We can thus expect that the shape of age-specific mortality rates for females in the early to mid-nineteenth century varied significantly from the shape of age-specific rates for females in the 1900–1902 DRA, even if the slope of the age profile is adjusted to account for suspected higher infant and

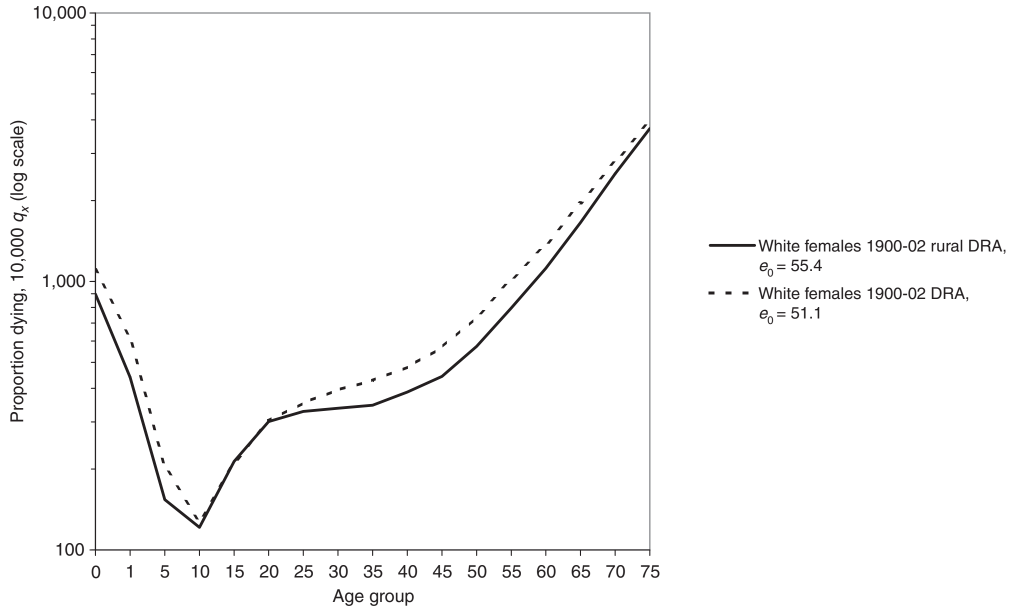

FIGURE 2. Proportion of dying by age group, white females in 1900–1902 death registration area.

child mortality relative to adult mortality prior to the onset of mortality decline.

Some indication of the possible bias can be seen in figure 2, which compares the age-specific mortality rates for white females in the 1900–1902 rural DRA life table with white females in the 1900–1902 overall DRA life table. Age-specific rates for white females in the rural DRA were noticeably lower than that for white females in the overall DRA at most ages, reflecting the overall higher life expectancy for females in rural areas. Mortality rates were roughly equal at ages 10–14, 20–24, and 25–29, however, and higher for rural females at age 15–19.[21] Although we cannot be sure of the causes, higher mortality in rural areas during adolescence and early adulthood are suggestive of higher death rates from tuberculosis and maternal mortality (Preston 1976; Henry 1989). Females age 20–49 residing in rural areas of the DRA had 9.4 percent more own children in the household than females in the overall DRA, increasing their exposure to maternal mortality and risk of contracting tuberculosis and other infectious diseases.

Among the four Princeton regional models, age-specific death rates for white males and females in the rural and overall 1900–1902 DRA had the closest correspondence with model west (after age 20, males in rural areas of the DRA had a closer relationship with the north model). Relative to the model west level corresponding to the same life expectancy at birth, however, female death rates in the 1900–1902 DRA and rural areas of the 1900–1902 DRA were much higher in peak childbearing years. The difference, as shown in figure 3, was especially pronounced for females residing in rural areas. With the exception of age groups between 15 and 35, there is remarkably close correspondence between west model level 15.17 and the mortality of women in the rural DRA. Age-specific death rates for rural females between ages 15 and 29, however, exceeded the level expected in model West by approximately 27 percent. The greatest divergence from the model pattern, 36 percent, was at ages 20–24. Although a similar pattern exists for males (not shown)—higher death rates at ages 5–34 for white males residing in rural areas of the 1900–1902 DRA relative to the corresponding model West level, lower rates at ages 40 and above—the differences were much smaller.

Similar "humps" in age-specific mortality rates for females between the approximate ages of 15 and 45 have been observed in other historical populations, including eighteenth- and early nineteenth-century American populations (Rutman and Rutman 1976; Logue 1991; Hacker 1996), the mostly rural eighteenth- and nineteenth-century populations studied by the Eurasia project (Alter, Manfredini, and Nystedt 2004), and the mid-nineteenth-century population of England (Wrigley and Schofield 1981/1989, 708–9). In their reconstruction of English population history, for example, Edward A. Wrigley and Roger S. Schofield noted that while age-specific mortality rates of males in the third English life table (1838–54) corresponded well with model North of the Princeton regional life tables, females had higher than expected rates from age 10 through age 35. The deviation from the model pattern prompted Wrigley and Schofield to construct their own model, based in part of the English life table and in part on the north model.

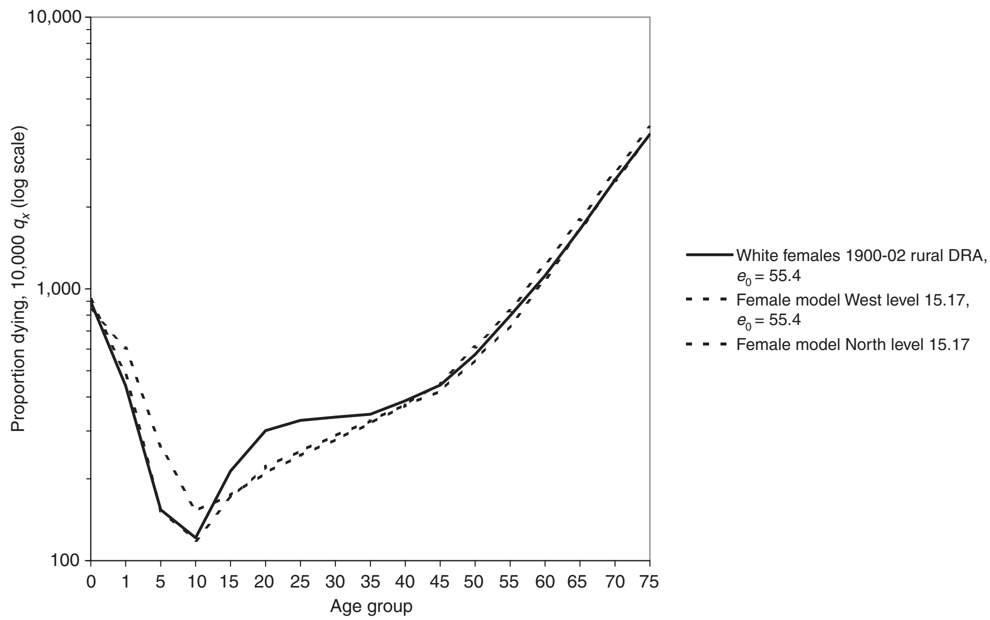

FIGURE 3. Proportion dying by age group, white females in rural areas of the 1900–1902 death registration area compared with Princeton model West and model North. q_x = proportion dying in age interval.

Did the same distinctive hump shape during childbearing years that characterized age-specific mortality rates for females in rural areas of the 1900–1902 DRA and various European and Asian populations also characterize the overall population of the nineteenth-century United States? Figure 4 shows the implied proportion dying in each age group from the Preston-Bennett 1850–60 life table shown in table 3 and the model West level corresponding to the equivalent life expectancy at age 10. Although the age pattern of mortality suggested by the Preston-Bennett life table is somewhat erratic, the distinctive deviation in age-specific mortality rates from the expected pattern is again evident. For females, the implied proportion dying in prime childbearing age groups 25–29 and 30–34 exceeded the implied proportion dying in age groups 35–39, 40–44, and 45–49. Although much less pronounced, a hump is also evident in the age-specific mortality pattern for white males. The two age profiles suggest the known age and sex profiles of tuberculosis mortality. The less pronounced hump for males may also indicate the absence of maternal mortality or different patterns of census coverage errors by age. Whatever the ultimate cause, the results of the Preston-Bennett life table suggest that the age-sex pattern of mortality in the nineteenth-century United States more closely resembled the pattern in the rural areas of the 1900–1902 DRA than the pattern in the overall DRA.

Another way of approaching the question is through examination of the sex mortality ratios by age. Despite higher life expectancies in the rural 1900–1902 DRA than in the overall DRA, the ratio of male-to-female mortality was lower at most ages in the rural DRA. The difference was especially pronounced during childbearing age.[22] White females in the rural DRA experienced excess mortality relative to males between age groups 15–19 and 40–44. In contrast, females in the overall DRA experienced lower mortality than males in all age groups. Among the nineteenth-century studies reporting lower female life expectancies in early adulthood cited above, most show excess females mortality relative to males in prime childbearing years. Alter, Manfredini, and Nystedt, for example, report excess female mortality from age 25 to 50 in six of the seven study populations in Sweden, Belgium, Italy, China, and Japan. In the rural village of Sart, Belgium, to cite a typical example, the ratio of male-to-female probability of dying in the interval 25–50 was 0.78 (2004, 334). England's third life table (1838–54) shows excess female mortality in all five-year age groups between age 10 and 40 (Wrigley and Schofield 1981/1989, 709), although the female disadvantage was modest. The lowest male-to-female mortality ratio, 0.95, was for the 25–29 age group. Excess female mortality was much higher in England's "healthy districts" (Woods 2000, 187), however, echoing the similar contrast between sex mortality ratios in the rural and overall 1900–1902 DRAs of the United States.

Although we lack death-registration data for the nineteenth-century United States, the 1860–1900 censuses of mortality allow the construction of sex differentials by age. Condran and Crimmins's (1979) analysis of these data indicated that, although the mortality censuses undercounted infant and elderly deaths, the *relative* undercount of males

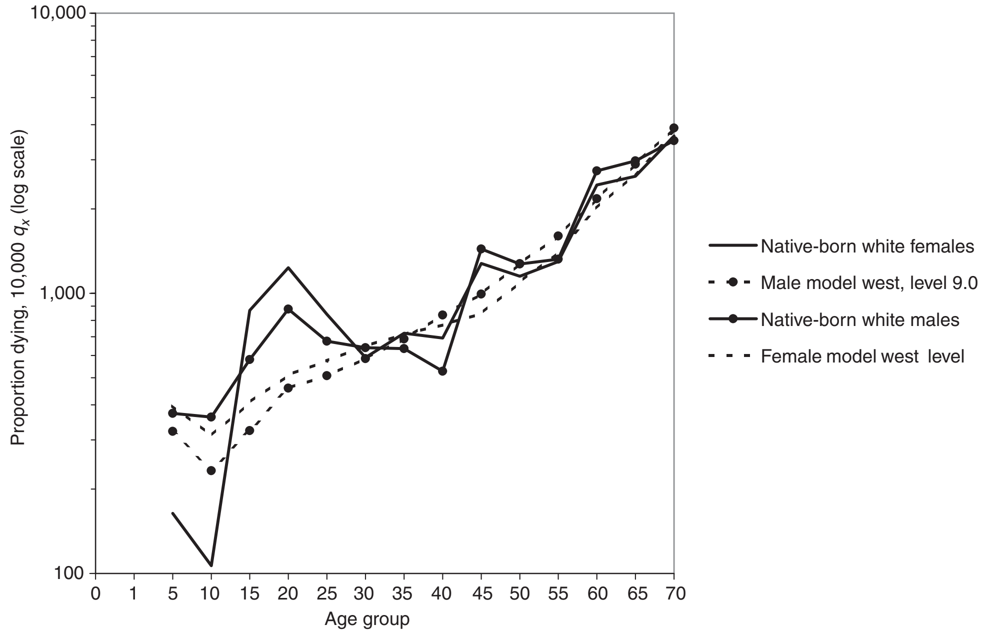

FIGURE 4. Proportion dying by age group, native-born whites in Preston-Bennett 1850–60 life table. q_x = proportion dying in age interval.

and females varied little by age. Figure 5 shows the average sex ratio in mortality in the 1860–80 censuses by age compared to the ratios in the overall 1900–1902 DRA and the rural areas of the 1900–1902 DRA. Figure 5 also includes a plot of the average sex mortality ratios in Haines's (1998) 1850–80 U.S. model life tables. Sex mortality ratios indicated by the census data suggest a similar pattern to the 1900–1902 rural DRA pattern: excess female mortality from adolescence through prime childbearing years and excess male mortality at other ages. Sex mortality ratios in Haines's life tables,

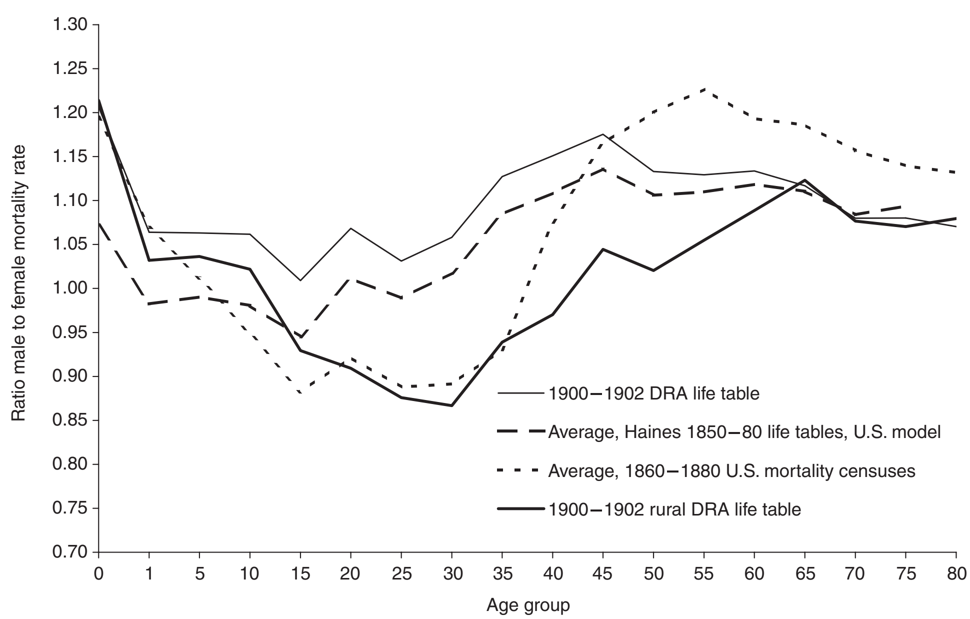

FIGURE 5. Ratio of male to female probability of dying by age (q_x). DRA = death registration area; q_x = proportion dying in age interval.

however, more closely conform to the 1900–1902 overall DRA. Although Haines's tables indicate modest excess female mortality in childhood and approximately equal sex ratios during prime childbearing years 20–34, the age pattern of sex mortality ratios is much closer to the overall 1900–1902 DRA pattern than to the rural DRA pattern. Haines's tables also suggest a lower sex differential in mortality in infancy than either the 1900–1902 overall or 1900–1902 rural DRA life tables.

Figures 4 and 5 strongly suggest that the 1900–1902 rural DRA life table is more representative of the shape of mortality in the nineteenth-century United States than the overall DRA life table. Age-specific mortality rates implied by the Preston-Bennett 1850–60 life table and sex mortality ratios by age in the 1860–80 censuses of mortality more closely conform to the pattern in the 1900–1902 rural DRA life table than the overall DRA life table (which was itself a closer match than model West). The correspondence should not be surprising: like the rural DRA life table, the population of the nineteenth-century United States was less urban, was more agricultural and had higher fertility than the population of the 1900–1902 DRA and populations used in the construction of model West. Although we cannot be certain of the true shape of age-specific mortality rates in the nineteenth century, the available evidence indicates that any model used to construct nineteenth-century life tables, especially life tables for the earlier part of the century, should draw more heavily from the 1900–1902 rural DRA life table than from the overall DRA life table.

New Decennial Life Tables, 1790–1910

Two life tables constructed by Glover (1921) for the 1900–1902 DRA are essential for this project: (1) the life table for the white population residing in the 10 DRA states and the District of Columbia, and (2) the life table for the white population in the rural areas of the DRA. When the life tables were published in 1921, the Census Bureau's definition of "urban" was considered cities of 8,000 or more inhabitants. All other places were considered rural. The Census Bureau subsequently redefined urban as places of 2,500 or more inhabitants. So although nominally nonurban, the 1900–1902 "rural" DRA life table is based in part on a population residing in the modern definition of an urban area, albeit modest towns and cities of 2,500 to 8,000 inhabitants.

As shown in table 2, the population living in the 1900–1902 DRA was predominately urban: over 60 percent lived in the modern definition of an urban area. Over 13 percent of the population in the rural areas of the DRA also lived in an urban area. The DRA covered 26.2 percent of the national 1900 population; the rural parts of the DRA only 12.0 percent.

What can be inferred about the level and pattern of national mortality in 1900–1902 given that nearly three-quarters of the population lived in states that were not part of the DRA? Although we could assume that the larger, more inclusive life table for the overall DRA is more representative of the national population, we know that urbanization, industrialization, nativity, and fertility in the DRA were not representative of the national population and likely affected the shape, level, and sex differential in mortality. A better choice might be the 1900–1902 rural DRA life table. Although a subset of the overall DRA, the rural population was more representative of the national population in terms of fertility, nativity, and occupation structure. Unsurprisingly, however, urbanization was higher in the nation as a whole than in the rural areas of the DRA and was likely the most important factor influencing mortality.

The simplest and most defensible inference is to combine the overall and rural DRA life tables, using appropriate weights to produce a life table reflecting the rate of urbanization in the nation as a whole. If we assume that that national population in 1901 was 40.2 percent urban (an interpolation of the Census Bureau's estimate of urbanization in the nation as a whole in 1900 and 1910), it is a simple matter to calculate the weight needed for each DRA life table and to combine the two to produce one life table representative of the nation's urban population.[23] Relative to the overall DRA life table, the resulting combined life table would increase estimates of white life expectancy at age 20 by 1.5 years for white males and 0.9 years for white females. Sex differentials in life expectancy at age 20 would fall from a 1.6-year female advantage in the overall DRA life table to a 0.9-year female advantage in the combined table.

The combined table could in turn be used as a model for earlier years: logits of the table's l_x values could be taken and new life tables generated by varying the slope and intercept shown in equation 2 to construct a predicted set of logits, calculating the l_x values by taking the anti-logits using equation 3, and constructing a new life table from the predicted l_x values.

There are several problems in such an approach. Most obviously, urbanization was increasing rapidly in the decades before the 1900 census. By design, the combined 1900–1902 life table is representative of urbanization in the 1900–1902 national population; nineteenth-century populations were far more rural. Haines's (1979) method is one possible way around this problem. Drawing on his analysis of available late nineteenth-century city and state life tables, Haines observed that the slope of age-specific mortality rates varied across time and by level of urbanization in a predictable way. Haines was thus able to set the slope of his model as needed to fit the period and level of urbanization.

Although a useful innovation, Haines's method cannot be applied uncritically to decades early in the nineteenth century. Most of the change observed in the slope of mortality likely reflected the impact of public health initiatives between 1880 and 1900 in the nation's largest cities, particularly efforts to clean water and milk supplies.[24] The net result was falling infant and early childhood mortality relative to adult mortality in large urban areas, despite rapidly increasing urbanization. Because most small cities made only modest attempts at public health initiatives before 1900 (Duffy 1990, chap. 12), it is much less certain if infant and childhood mortality fell relative to adult mortality for the nation as a whole between 1850 and 1900, which Haines's model predicts. Indeed, as Haines (1979, 300–301) noted, the Princeton west model suggests the opposite. Between levels 9 and 13—equivalent to an increase in female life expectancy from 40 to 50 years and roughly spanning the increase in life expectancy in the late nineteenth-century United States—the west model suggests that infant and childhood mortality should increase relative to adult mortality. Only at mortality levels above level 13 does infant and childhood mortality begin to decline faster than adult mortality.

Given this uncertainty, a better approach would be to create a unique standard for each decade of the nineteenth century by repeating the weighting exercise of the 1900–1902 DRA and 1900–1902 rural DRA life tables described above, using the appropriate weights to yield a new standard life table representative of the urbanization level in each decade. Table 7 shows the results of that exercise. Included in the table are estimates of the mid-census level of urbanization in each decade (an average of the percentage urban in each of the beginning and ending censuses), the corresponding proportional weights of the 1900–1902 overall and rural DRA life tables used to create each standard, and the resulting logits of the tables' l_x values by age and sex. Before 1850, the national level of urbanization was below that estimated in the rural 1900–1902 DRA table. It was therefore assumed that the rural 1900–1902 table represented the standard mortality pattern for all decades before to 1850. After 1850 urbanization began to exceed the level of urbanization in the 1900–1902 rural DRA life table, requiring increasing weight to be given to the overall DRA life table. The applied weighting of the 1900–1902 overall DRA life table increased from 0.09 in the 1850–59 decade to 0.51 in the 1890–99 decade.

From there it was a simple matter of varying the intercept in equation 2 and constructing a new life table to fit the estimates of adult life expectancy shown in tables 4 and 5. With one exception, the resulting life tables are shown in table 8. The exception is the 1860–69 life table for white males, which was modified to account for high mortality among males of military age during the Civil War. It was constructed in three steps. First, a "base" life table for the 1860–69 period was constructed by using the average of the 1850–59 and 1870–79 estimates of male life expectancy at age 20. Second, an estimate of excess male deaths in the 1860–70 intercensal period was made by cohort using two-census survival methods.[25] Finally, the excess male deaths were added to the base life table. Table 9 shows the results for each year of the war.[26] Unsurprisingly, mortality was highest in 1864, the last full year of the conflict. The estimates imply a white male life expectancy at birth of 25.9 years, likely the lowest level in U.S. history. Although based on crude estimates, the method retains the unusual risk of early death among young white males in the war. The resulting life table for the 1860s suggests a male life expectancy at age 20 of 35.1 years, approximately two years lower than the adjusted Kunze (1979) and Pope (1992) estimate. Although the base life table and the number of excess male deaths could be adjusted to yield a perfect match, it is unclear which estimate to adjust. It is also possible that the genealogical samples, which are known to underrepresent individuals who do not marry or reproduce, are biased against soldiers participating and dying in the war. It was therefore decided to make no further adjustments to the life table.

Figures 6, 7, and 8 compare some of the new life table estimates with Haines's (1998) life table estimates. As shown in figure 6, the new life tables describe a decline in life expectancy at birth from approximately 44 years in the late eighteenth century to just over 40 years in the 1840s. Although the models assumed a slight male advantage in life expectancy at age 20, higher male mortality in infancy pushed female life expectancy at birth slightly above the male estimate. Life expectancy at birth then declined another 3–4 years in the 1850s to approximately 37 years. The decline is largely the result of the model's prediction of increased infant mortality. Although the decline in adult life expectancy between the 1840s and 1850s was relatively modest (1.4 years), the model suggests that infant mortality rates rose from 215 to 247 per 1,000 for white males and from 190 to 222 for white females. Life expectancy reached an even lower level in the 1860s for white males—the result of the U.S. Civil War—but then increased rapidly with estimates for white females for the remainder of the century. Life expectancy for white females increased more rapidly. By the 1890s, white females enjoyed about a two-year advantage in life expectancy at birth.

Haines's (1998) estimates are plotted with a marker to emphasize their limitation to individual census years. In general, Haines's life tables document a similar pattern of low life expectancy at midcentury and a rapid increase late in the century. Haines's estimates for 1860 are relatively high, however, whereas his estimate for the 1880s is relatively low. It is difficult to know what to make of the differences. The substantial decline in life expectancy between 1870 in 1880, in particular, does not correspond with known epidemics or the qualitative literature on the mortality decline in the United States. The decline may reflect that 1880 was a particularly unhealthy year or be an artifact of differential census enumeration. The 1870 census has been long suspected to have undercounted the population and may well have undercounted mortality as well. The 1880 census, on the other hand, benefitted from a shift from enumeration by United States marshals to enumeration by trained enumerators, a sharp increase in the number of enumerators relative to the population, and the supplementation of mortality data in the census with available death registration data.

Figures 7 and 8 compare the proportions dying in five-year age intervals in male and female life tables selected

TABLE 7. Standard Life Table Logit Values, $Y_s(x)$, for Decennial Life Tables

	Decade										
Age	1790–99	1800–09	1810–19	1820–29	1830–39	1840–49	1850–59	1860–69	1870–79	1880–89	1890–99
Males											
0											
1	−1.0505	−1.0505	−1.0505	−1.0505	−1.0505	−1.0505	−1.0390	−1.0254	−1.0147	−1.0029	−0.9888
2	−0.9476	−0.9476	−0.9476	−0.9476	−0.9476	−0.9476	−0.9344	−0.9188	−0.9066	−0.8932	−0.8772
3	−0.9081	−0.9081	−0.9081	−0.9081	−0.9081	−0.9081	−0.8941	−0.8775	−0.8646	−0.8504	−0.8334
4	−0.8840	−0.8840	−0.8840	−0.8840	−0.8840	−0.8840	−0.8694	−0.8523	−0.8389	−0.8243	−0.8068
5	−0.8656	−0.8656	−0.8656	−0.8656	−0.8656	−0.8656	−0.8508	−0.8334	−0.8198	−0.8049	−0.7871
10	−0.8144	−0.8144	−0.8144	−0.8144	−0.8144	−0.8144	−0.7992	−0.7813	−0.7673	−0.7521	−0.7338
15	−0.7775	−0.7775	−0.7775	−0.7775	−0.7775	−0.7775	−0.7629	−0.7457	−0.7322	−0.7175	−0.6999
20	−0.7225	−0.7225	−0.7225	−0.7225	−0.7225	−0.7225	−0.7088	−0.6926	−0.6799	−0.6660	−0.6494
25	−0.6538	−0.6538	−0.6538	−0.6538	−0.6538	−0.6538	−0.6403	−0.6246	−0.6121	−0.5985	−0.5822
30	−0.5887	−0.5887	−0.5887	−0.5887	−0.5887	−0.5887	−0.5750	−0.5589	−0.5463	−0.5324	−0.5158
35	−0.5285	−0.5285	−0.5285	−0.5285	−0.5285	−0.5285	−0.5137	−0.4964	−0.4827	−0.4678	−0.4499
40	−0.4673	−0.4673	−0.4673	−0.4673	−0.4673	−0.4673	−0.4512	−0.4322	−0.4173	−0.4011	−0.3816
45	−0.4024	−0.4024	−0.4024	−0.4024	−0.4024	−0.4024	−0.3850	−0.3645	−0.3485	−0.3310	−0.3100
50	−0.3296	−0.3296	−0.3296	−0.3296	−0.3296	−0.3296	−0.3109	−0.2889	−0.2718	−0.2530	−0.2305
55	−0.2458	−0.2458	−0.2458	−0.2458	−0.2458	−0.2458	−0.2258	−0.2023	−0.1839	−0.1638	−0.1398
60	−0.1375	−0.1375	−0.1375	−0.1375	−0.1375	−0.1375	−0.1163	−0.0914	−0.0718	−0.0505	−0.0249
65	0.0019	0.0019	0.0019	0.0019	0.0019	0.0019	0.0237	0.0494	0.0696	0.0917	0.1182
70	0.1897	0.1897	0.1897	0.1897	0.1897	0.1897	0.2114	0.2370	0.2572	0.2794	0.3062
75	0.4329	0.4329	0.4329	0.4329	0.4329	0.4329	0.4541	0.4792	0.4991	0.5211	0.5477

(Continued on next page)

TABLE 7. Standard Life Table Logit Values, $Y_s(x)$, for Decennial Life Tables (*Continued*)

	Decade										
Age	1790–99	1800–09	1810–19	1820–29	1830–39	1840–49	1850–59	1860–69	1870–79	1880–89	1890–99
Females											
0											
1	−1.1581	−1.1581	−1.1581	−1.1581	−1.1581	−1.1581	−1.1465	−1.1327	−1.1219	−1.1101	−1.0959
2	−1.0471	−1.0471	−1.0471	−1.0471	−1.0471	−1.0471	−1.0336	−1.0177	−1.0053	−0.9917	−0.9755
3	−1.0004	−1.0004	−1.0004	−1.0004	−1.0004	−1.0004	−0.9864	−0.9699	−0.9570	−0.9429	−0.9260
4	−0.9726	−0.9726	−0.9726	−0.9726	−0.9726	−0.9726	−0.9582	−0.9412	−0.9279	−0.9134	−0.8961
5	−0.9512	−0.9512	−0.9512	−0.9512	−0.9512	−0.9512	−0.9365	−0.9193	−0.9058	−0.8911	−0.8736
10	−0.8944	−0.8944	−0.8944	−0.8944	−0.8944	−0.8944	−0.8794	−0.8619	−0.8482	−0.8333	−0.8154
15	−0.8532	−0.8532	−0.8532	−0.8532	−0.8532	−0.8532	−0.8391	−0.8225	−0.8095	−0.7953	−0.7783
20	−0.7867	−0.7867	−0.7867	−0.7867	−0.7867	−0.7867	−0.7742	−0.7593	−0.7477	−0.7349	−0.7196
25	−0.7037	−0.7037	−0.7037	−0.7037	−0.7037	−0.7037	−0.6927	−0.6796	−0.6693	−0.6580	−0.6444
30	−0.6241	−0.6241	−0.6241	−0.6241	−0.6241	−0.6241	−0.6138	−0.6017	−0.5921	−0.5815	−0.5687
35	−0.5514	−0.5514	−0.5514	−0.5514	−0.5514	−0.5514	−0.5411	−0.5289	−0.5193	−0.5087	−0.4959
40	−0.4842	−0.4842	−0.4842	−0.4842	−0.4842	−0.4842	−0.4735	−0.4608	−0.4507	−0.4397	−0.4264
45	−0.4159	−0.4159	−0.4159	−0.4159	−0.4159	−0.4159	−0.4047	−0.3914	−0.3810	−0.3694	−0.3555
50	−0.3448	−0.3448	−0.3448	−0.3448	−0.3448	−0.3448	−0.3328	−0.3187	−0.3075	−0.2952	−0.2804
55	−0.2612	−0.2612	−0.2612	−0.2612	−0.2612	−0.2612	−0.2483	−0.2330	−0.2210	−0.2078	−0.1918
60	−0.1566	−0.1566	−0.1566	−0.1566	−0.1566	−0.1566	−0.1426	−0.1261	−0.1130	−0.0986	−0.0813
65	−0.0259	−0.0259	−0.0259	−0.0259	−0.0259	−0.0259	−0.0111	0.0064	0.0203	0.0355	0.0539
70	0.1449	0.1449	0.1449	0.1449	0.1449	0.1449	0.1605	0.1791	0.1938	0.2100	0.2295
75	0.3763	0.3763	0.3763	0.3763	0.3763	0.3763	0.3920	0.4107	0.4256	0.4421	0.4620
80+	0.6887	0.6887	0.6887	0.6887	0.6887	0.6887	0.7043	0.7230	0.7379	0.7544	0.7746
% urban	5.6	6.7	7.2	8.0	9.8	13.0	17.5	22.7	26.9	31.6	37.3
1901 DRA weight	0.00	0.00	0.00	0.00	0.00	0.00	0.09	0.20	0.29	0.39	0.51
1901 rural DRA weight	1.00	1.00	1.00	1.00	1.00	1.00	0.91	0.80	0.71	0.61	0.49

TABLE 8. New Life Tables for the White Population of the United States, 1780–1900

	White males, 1790–99							White females, 1790–99					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.1797	100000	17969	87961	4409061	44.1	0	0.1603	100000	16030	89580	4418888	44.2
1	0.0394	82031	3235	80122	4321100	52.7	1	0.0383	83970	3219	82070	4329308	51.6
2	0.0171	78796	1349	78081	4240978	53.8	2	0.0185	80751	1491	79960	4247237	52.6
3	0.0110	77447	854	77003	4162897	53.8	3	0.0117	79260	929	78777	4167277	52.6
4	0.0087	76593	666	76246	4085894	53.3	4	0.0094	78331	735	77948	4088500	52.2
5	0.0253	75927	1921	374832	4009648	52.8	5	0.0263	77595	2039	382878	4010552	51.7
10	0.0195	74006	1445	366418	3634815	49.1	10	0.0205	75556	1552	373900	3627674	48.0
15	0.0309	72561	2243	357200	3268397	45.0	15	0.0356	74004	2637	363427	3253774	44.0
20	0.0419	70319	2949	344222	2911197	41.4	20	0.0492	71367	3510	348059	2890347	40.5
25	0.0434	67370	2924	329541	2566974	38.1	25	0.0526	67857	3567	330367	2542289	37.5
30	0.0435	64446	2802	315226	2237433	34.7	30	0.0529	64290	3404	312942	2211921	34.4
35	0.0475	61644	2931	300892	1922207	31.2	35	0.0532	60887	3242	296329	1898979	31.2
40	0.0541	58713	3177	285621	1621316	27.6	40	0.0584	57645	3368	279805	1602650	27.8
45	0.0652	55536	3622	268624	1335695	24.1	45	0.0653	54277	3543	262528	1322845	24.4
50	0.0806	51914	4185	249109	1067071	20.6	50	0.0823	50734	4177	243230	1060317	20.9
55	0.1122	47729	5356	225256	817962	17.1	55	0.1106	46558	5147	219921	817087	17.5
60	0.1564	42373	6627	195298	592706	14.0	60	0.1489	41411	6167	191636	597166	14.4
65	0.2266	35746	8099	158483	397407	11.1	65	0.2087	35244	7357	157827	405529	11.5
70	0.3119	27647	8622	116679	238925	8.6	70	0.2979	27887	8307	118667	247703	8.9
75	0.4312	19025	8203	74616	122245	6.4	75	0.4111	19580	8050	77776	129036	6.6
80+	1.0000	10822	10822	47629	47629	4.4	80+	1.0000	11530	11530	51260	51260	4.4

	White males, 1800–09							White females, 1800–09					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.2036	100000	20357	86361	4138061	41.4	0	0.1807	100000	18070	88254	4166944	41.7
1	0.0444	79643	3540	77554	4051700	50.9	1	0.0430	81930	3523	79851	4078690	49.8
2	0.0193	76103	1465	75327	3974146	52.2	2	0.0207	78406	1620	77547	3998839	51.0
3	0.0124	74638	925	74157	3898819	52.2	3	0.0131	76786	1006	76263	3921292	51.1
4	0.0098	73713	719	73340	3824662	51.9	4	0.0105	75780	794	75367	3845029	50.7
5	0.0283	72994	2065	359809	3751323	51.4	5	0.0293	74986	2194	369444	3769662	50.3
10	0.0218	70929	1545	350784	3391514	47.8	10	0.0228	72792	1660	359809	3400218	46.7
15	0.0344	69384	2384	340961	3040730	43.8	15	0.0394	71132	2804	348648	3040409	42.7
20	0.0464	67000	3109	327228	2699769	40.3	20	0.0541	68327	3698	332393	2691761	39.4
25	0.0478	63891	3054	311820	2372541	37.1	25	0.0575	64630	3718	313853	2359368	36.5
30	0.0477	60837	2901	296931	2060721	33.9	30	0.0577	60911	3512	295776	2045515	33.6
35	0.0519	57936	3007	282159	1763791	30.4	35	0.0577	57399	3313	278714	1749739	30.5
40	0.0588	54928	3229	266569	1481631	27.0	40	0.0630	54086	3409	261910	1471025	27.2
45	0.0704	51699	3642	249392	1215062	23.5	45	0.0700	50678	3550	244513	1209115	23.9
50	0.0865	48058	4157	229894	965670	20.1	50	0.0878	47128	4139	225292	964602	20.5
55	0.1195	43900	5244	206389	735776	16.8	55	0.1171	42989	5033	202363	739310	17.2
60	0.1648	38656	6371	177350	529387	13.7	60	0.1563	37956	5934	174947	536947	14.1
65	0.2359	32284	7616	142382	352036	10.9	65	0.2169	32023	6945	142751	362000	11.3
70	0.3206	24668	7908	103571	209654	8.5	70	0.3059	25078	7672	106208	219249	8.7
75	0.4380	16760	7340	65449	106083	6.3	75	0.4176	17406	7268	68857	113041	6.5
80+	1.0000	9420	9420	40634	40634	4.3	80+	1.0000	10137	10137	44184	44184	4.4

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TABLE 8. New Life Tables for the White Population of the United States, 1780–1900 (*Continued*)

	White males, 1810–19							White females, 1810–19					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.2190	100000	21898	85328	3975007	39.8	0	0.1938	100000	19384	87400	4015840	40.2
1	0.0476	78102	3721	75906	3889679	49.8	1	0.0460	80616	3707	78428	3928440	48.7
2	0.0206	74380	1533	73568	3813773	51.3	2	0.0221	76908	1697	76009	3850011	50.1
3	0.0133	72847	965	72345	3740206	51.3	3	0.0140	75211	1051	74665	3774002	50.2
4	0.0104	71882	749	71492	3667860	51.0	4	0.0112	74160	829	73729	3699338	49.9
5	0.0302	71133	2147	350295	3596368	50.6	5	0.0311	73331	2283	360950	3625608	49.4
10	0.0232	68985	1601	340925	3246073	47.1	10	0.0242	71049	1722	350938	3264659	45.9
15	0.0365	67385	2461	330770	2905148	43.1	15	0.0418	69327	2897	339392	2913720	42.0
20	0.0492	64923	3193	316635	2574378	39.7	20	0.0572	66430	3798	322655	2574329	38.8
25	0.0505	61731	3119	300856	2257743	36.6	25	0.0606	62632	3795	303673	2251674	36.0
30	0.0503	58612	2946	285694	1956887	33.4	30	0.0605	58837	3562	285282	1948001	33.1
35	0.0546	55666	3037	270737	1671192	30.0	35	0.0604	55275	3339	268029	1662719	30.1
40	0.0616	52629	3242	255041	1400455	26.6	40	0.0658	51936	3417	251138	1394690	26.9
45	0.0736	49387	3633	237853	1145414	23.2	45	0.0729	48519	3536	233756	1143552	23.6
50	0.0900	45754	4118	218475	907561	19.8	50	0.0911	44983	4096	214674	909796	20.2
55	0.1237	41636	5150	195306	689085	16.6	55	0.1209	40887	4942	192079	695122	17.0
60	0.1697	36486	6190	166956	493780	13.5	60	0.1606	35945	5772	165295	503043	14.0
65	0.2412	30296	7306	133214	326824	10.8	65	0.2215	30173	6682	134158	337749	11.2
70	0.3254	22990	7481	96246	193610	8.4	70	0.3104	23490	7292	99223	203590	8.7
75	0.4416	15509	6849	60420	97363	6.3	75	0.4211	16199	6822	63941	104367	6.4
80+	1.0000	8659	8659	36943	36943	4.3	80+	1.0000	9377	9377	40426	40426	4.3

	White males, 1820–29							White females, 1820–29					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.2166	100000	21663	85486	3999382	40.0	0	0.1918	100000	19184	87531	4038406	40.4
1	0.0472	78337	3694	76157	3913896	50.0	1	0.0455	80816	3680	78645	3950875	48.9
2	0.0204	74643	1523	73836	3837739	51.4	2	0.0219	77137	1686	76243	3872230	50.2
3	0.0131	73120	959	72621	3763903	51.5	3	0.0138	75451	1045	74908	3795986	50.3
4	0.0103	72160	745	71773	3691282	51.2	4	0.0111	74406	824	73978	3721079	50.0
5	0.0299	71416	2135	351740	3619509	50.7	5	0.0308	73583	2269	362239	3647101	49.6
10	0.0230	69280	1593	342420	3267769	47.2	10	0.0240	71313	1713	352283	3284862	46.1
15	0.0362	67688	2450	332314	2925349	43.2	15	0.0414	69600	2883	340793	2932578	42.1
20	0.0488	65238	3181	318236	2593035	39.7	20	0.0567	66717	3784	324125	2591785	38.8
25	0.0501	62057	3110	302509	2274799	36.7	25	0.0601	62933	3784	305206	2267660	36.0
30	0.0499	58947	2940	287384	1972290	33.5	30	0.0601	59149	3555	286858	1962454	33.2
35	0.0542	56007	3033	272451	1684906	30.1	35	0.0600	55594	3336	269629	1675597	30.1
40	0.0612	52974	3241	256766	1412455	26.7	40	0.0654	52258	3416	252748	1405967	26.9
45	0.0731	49733	3636	239576	1155688	23.2	45	0.0725	48841	3539	235359	1153219	23.6
50	0.0895	46097	4125	220175	916113	19.9	50	0.0906	45302	4104	216252	917860	20.3
55	0.1231	41972	5165	196949	695938	16.6	55	0.1203	41199	4956	193603	701608	17.0
60	0.1689	36807	6218	168490	498989	13.6	60	0.1600	36242	5797	166719	508005	14.0
65	0.2404	30589	7353	134560	330499	10.8	65	0.2208	30445	6722	135420	341286	11.2
70	0.3247	23235	7545	97315	195938	8.4	70	0.3098	23723	7348	100244	205866	8.7
75	0.4411	15691	6921	61150	98623	6.3	75	0.4206	16375	6887	64655	105622	6.5
80+	1.0000	8769	8769	37473	37473	4.3	80+	1.0000	9487	9487	40966	40966	4.3

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TABLE 8. New Life Tables for the White Population of the United States, 1780–1900 (*Continued*)

	White males, 1830–39							White females, 1830–39					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.2119	100000	21188	85804	4049121	40.5	0	0.1878	100000	18779	87794	4084478	40.8
1	0.0462	78812	3639	76665	3963317	50.3	1	0.0446	81221	3624	79083	3996684	49.2
2	0.0200	75173	1502	74377	3886652	51.7	2	0.0214	77597	1662	76716	3917601	50.5
3	0.0129	73671	947	73178	3812275	51.7	3	0.0136	75935	1031	75399	3840884	50.6
4	0.0101	72724	736	72341	3739097	51.4	4	0.0109	74904	813	74481	3765485	50.3
5	0.0293	71988	2111	354664	3666756	50.9	5	0.0303	74091	2243	364849	3691004	49.8
10	0.0226	69878	1576	345448	3312092	47.4	10	0.0236	71849	1694	355007	3326154	46.3
15	0.0355	68302	2427	335441	2966644	43.4	15	0.0407	70154	2855	343633	2971147	42.4
20	0.0479	65875	3156	321484	2631202	39.9	20	0.0558	67299	3754	327110	2627514	39.0
25	0.0493	62719	3091	305867	2309718	36.8	25	0.0592	63545	3761	308322	2300404	36.2
30	0.0491	59628	2927	290822	2003852	33.6	30	0.0592	59784	3541	290066	1992083	33.3
35	0.0533	56701	3025	275942	1713030	30.2	35	0.0592	56243	3329	272892	1702017	30.3
40	0.0603	53676	3238	260285	1437089	26.8	40	0.0645	52914	3415	256033	1429124	27.0
45	0.0721	50438	3639	243093	1176803	23.3	45	0.0716	49499	3544	238636	1173091	23.7
50	0.0884	46799	4138	223651	933710	20.0	50	0.0896	45955	4117	219483	934455	20.3
55	0.1218	42661	5195	200318	710059	16.6	55	0.1192	41838	4985	196726	714972	17.1
60	0.1675	37466	6275	171642	509742	13.6	60	0.1587	36853	5847	169646	518245	14.1
65	0.2388	31191	7449	137334	338099	10.8	65	0.2194	31005	6803	138020	348600	11.2
70	0.3233	23742	7675	99525	200765	8.5	70	0.3084	24203	7464	102352	210580	8.7
75	0.4400	16068	7070	62664	101240	6.3	75	0.4195	16738	7022	66136	108227	6.5
80+	1.0000	8998	8998	38577	38577	4.3	80+	1.0000	9716	9716	42092	42092	4.3

	White males, 1840–49							White females, 1840–49					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.2286	100000	22862	84682	3877227	38.8	0	0.2020	100000	20205	86867	3925388	39.3
1	0.0496	77138	3829	74879	3792544	49.2	1	0.0478	79795	3818	77543	3838521	48.1
2	0.0215	73309	1573	72475	3717666	50.7	2	0.0229	75978	1743	75054	3760978	49.5
3	0.0138	71736	989	71222	3645190	50.8	3	0.0145	74235	1078	73674	3685924	49.7
4	0.0108	70747	767	70348	3573969	50.5	4	0.0116	73157	849	72716	3612250	49.4
5	0.0314	69980	2194	344415	3503620	50.1	5	0.0323	72308	2334	355705	3539534	49.0
10	0.0241	67786	1632	334849	3159205	46.6	10	0.0251	69974	1757	345476	3183829	45.5
15	0.0379	66154	2504	324509	2824356	42.7	15	0.0432	68217	2949	333710	2838353	41.6
20	0.0509	63650	3238	310154	2499847	39.3	20	0.0590	65268	3853	316705	2504643	38.4
25	0.0522	60412	3152	294179	2189693	36.2	25	0.0624	61414	3835	297485	2187938	35.6
30	0.0518	57260	2967	278882	1895514	33.1	30	0.0623	57580	3586	278935	1890452	32.8
35	0.0562	54293	3049	263842	1616632	29.8	35	0.0620	53994	3350	261597	1611517	29.8
40	0.0633	51244	3243	248114	1352790	26.4	40	0.0674	50645	3415	244685	1349920	26.7
45	0.0754	48001	3621	230955	1104676	23.0	45	0.0746	47229	3522	227341	1105236	23.4
50	0.0921	44381	4086	211688	873721	19.7	50	0.0930	43707	4064	208376	877894	20.1
55	0.1262	40294	5084	188762	662033	16.4	55	0.1231	39643	4880	186017	669518	16.9
60	0.1725	35210	6073	160870	473271	13.4	60	0.1631	34764	5669	159646	483500	13.9
65	0.2442	29137	7115	127899	312402	10.7	65	0.2241	29095	6521	129173	323854	11.1
70	0.3282	22022	7227	92045	184503	8.4	70	0.3130	22574	7065	95209	194681	8.6
75	0.4437	14796	6565	57565	92458	6.2	75	0.4231	15509	6562	61141	99472	6.4
80+	1.0000	8231	8231	34893	34893	4.2	80+	1.0000	8947	8947	38331	38331	4.3

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TABLE 8. New Life Tables for the White Population of the United States, 1780–1900 (*Continued*)

	White males, 1850–59							White females, 1850–59					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.2465	100000	24647	83486	3684941	36.8	0	0.2220	100000	22198	85572	3713004	37.1
1	0.0542	75353	4087	72941	3601455	47.8	1	0.0532	77802	4141	75359	3627432	46.6
2	0.0236	71265	1679	70376	3528514	49.5	2	0.0254	73661	1871	72669	3552073	48.2
3	0.0151	69587	1052	69039	3458138	49.7	3	0.0161	71790	1158	71188	3479403	48.5
4	0.0118	68534	810	68113	3389099	49.5	4	0.0128	70632	905	70162	3408215	48.3
5	0.0339	67724	2296	332882	3320986	49.0	5	0.0353	69728	2464	342478	3338053	47.9
10	0.0254	65429	1660	322992	2988104	45.7	10	0.0268	67263	1801	331815	2995575	44.5
15	0.0397	63768	2534	312505	2665112	41.8	15	0.0457	65463	2991	319835	2663760	40.7
20	0.0538	61234	3295	297932	2352607	38.4	20	0.0623	62471	3892	302628	2343925	37.5
25	0.0555	57939	3213	281661	2054676	35.5	25	0.0661	58580	3870	283224	2041298	34.8
30	0.0557	54726	3050	266003	1773015	32.4	30	0.0662	54710	3623	264493	1758074	32.1
35	0.0605	51676	3127	250561	1507012	29.2	35	0.0661	51087	3379	246989	1493581	29.2
40	0.0679	48549	3295	234505	1256451	25.9	40	0.0716	47708	3417	229998	1246592	26.1
45	0.0805	45253	3641	217165	1021946	22.6	45	0.0792	44291	3510	212680	1016594	23.0
50	0.0977	41613	4066	197899	804781	19.3	50	0.0984	40781	4012	193875	803914	19.7
55	0.1326	37547	4980	175285	606882	16.2	55	0.1295	36769	4762	171937	610040	16.6
60	0.1790	32567	5828	148264	431597	13.3	60	0.1698	32006	5434	146446	438102	13.7
65	0.2502	26739	6690	116968	283333	10.6	65	0.2313	26572	6145	117499	291656	11.0
70	0.3331	20049	6678	83549	166365	8.3	70	0.3190	20427	6516	85846	174157	8.5
75	0.4476	13371	5985	51893	82816	6.2	75	0.4276	13911	5948	54686	88311	6.3
80+	1.0000	7386	7386	30923	30923	4.2	80+	1.0000	7963	7963	33625	33625	4.2

	White males, 1860–69							White females, 1860–69					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.2071	100000	20708	86125	3506917	35.1	0	0.1893	100000	18930	87696	4057822	40.6
1	0.0469	79292	3718	77098	3420791	43.1	1	0.0467	81070	3783	78838	3970126	49.0
2	0.0206	75573	1556	74749	3343693	44.2	2	0.0223	77287	1723	76374	3891288	50.3
3	0.0132	74017	980	73508	3268945	44.2	3	0.0142	75565	1076	75005	3814913	50.5
4	0.0103	73037	753	72645	3195437	43.8	4	0.0113	74489	840	74052	3739908	50.2
5	0.0295	72284	2133	356085	3122792	43.2	5	0.0311	73649	2288	362524	3665856	49.8
10	0.0216	70150	1513	346968	2766706	39.4	10	0.0230	71361	1638	352709	3303332	46.3
15	0.0812	68637	5572	329256	2419738	35.3	15	0.0392	69723	2731	341787	2950623	42.3
20	0.1049	63065	6613	298795	2090482	33.1	20	0.0540	66992	3616	325919	2608837	38.9
25	0.0999	56452	5639	268165	1791687	31.7	25	0.0582	63376	3687	307660	2282917	36.0
30	0.1041	50814	5291	240840	1523522	30.0	30	0.0594	59688	3545	289580	1975257	33.1
35	0.0927	45522	4222	217057	1282682	28.2	35	0.0602	56143	3378	272272	1685677	30.0
40	0.0831	41301	3432	197924	1065625	25.8	40	0.0657	52765	3465	255163	1413406	26.8
45	0.0819	37869	3103	181586	867701	22.9	45	0.0735	49300	3626	237436	1158242	23.5
50	0.0943	34766	3279	165630	686115	19.7	50	0.0922	45674	4209	217849	920806	20.2
55	0.1251	31487	3938	147589	520484	16.5	55	0.1225	41465	5081	194623	702957	17.0
60	0.1697	27549	4675	126057	372896	13.5	60	0.1617	36384	5885	167208	508334	14.0
65	0.2394	22874	5476	100678	246839	10.8	65	0.2228	30499	6796	135506	341126	11.2
70	0.3229	17397	5617	72943	146161	8.4	70	0.3101	23703	7351	100139	205620	8.7
75	0.4402	11780	5186	45936	73218	6.2	75	0.4205	16352	6876	64570	105481	6.5
80+	1.0000	6594	6594	27282	27282	4.1	80+	1.0000	9476	9476	40910	40910	4.3

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TABLE 8. New Life Tables for the White Population of the United States, 1780–1900 (*Continued*)

	White males, 1870–79							White females, 1870–79				
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.1753	100000	17533	88253	4398899	44.0	0	0.1616	100000	16164	89493	4387395	43.9
1	0.0406	82467	3348	80491	4310646	52.3	1	0.0407	83836	3415	81821	4297902	51.3
2	0.0180	79119	1422	78365	4230154	53.5	2	0.0195	80421	1566	79591	4216081	52.4
3	0.0116	77696	900	77228	4151790	53.4	3	0.0125	78855	987	78341	4136491	52.5
4	0.0090	76796	691	76436	4074562	53.1	4	0.0099	77868	769	77468	4058149	52.1
5	0.0257	76105	1958	375629	3998125	52.5	5	0.0272	77099	2099	380246	3980681	51.6
10	0.0185	74147	1369	367311	3622496	48.9	10	0.0197	75000	1480	371300	3600435	48.0
15	0.0291	72778	2120	358587	3255185	44.7	15	0.0337	73520	2475	361412	3229135	43.9
20	0.0409	70657	2888	346066	2896598	41.0	20	0.0468	71045	3327	346906	2867723	40.4
25	0.0434	67769	2941	331493	2550532	37.6	25	0.0512	67717	3465	329924	2520817	37.2
30	0.0455	64828	2949	316767	2219039	34.2	30	0.0530	64252	3408	312742	2190893	34.1
35	0.0506	61879	3129	301571	1902272	30.7	35	0.0544	60844	3310	295946	1878151	30.9
40	0.0574	58750	3373	285316	1600701	27.2	40	0.0598	57534	3441	279069	1582205	27.5
45	0.0689	55377	3816	267343	1315385	23.8	45	0.0677	54094	3663	261310	1303136	24.1
50	0.0851	51560	4387	246834	1048042	20.3	50	0.0856	50431	4317	241360	1041825	20.7
55	0.1172	47173	5529	222044	801208	17.0	55	0.1150	46113	5304	217307	800466	17.4
60	0.1602	41645	6672	191542	579164	13.9	60	0.1531	40810	6247	188431	583158	14.3
65	0.2284	34972	7989	154888	387622	11.1	65	0.2135	34563	7379	154365	394728	11.4
70	0.3124	26983	8429	113844	232735	8.6	70	0.3005	27183	8167	115499	240362	8.8
75	0.4323	18554	8021	72720	118891	6.4	75	0.4127	19016	7847	75462	124863	6.6
80+	1.0000	10533	10533	46171	46171	4.4	80+	1.0000	11169	11169	49401	49401	4.4
	White males, 1880–89							White females, 1880–89					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.1494	100000	14936	89993	4708224	47.1	0	0.1317	100000	13168	91441	4795214	48.0
1	0.0354	85064	3008	83290	4618231	54.3	1	0.0340	86832	2950	85091	4703773	54.2
2	0.0158	82056	1296	81370	4534941	55.3	2	0.0163	83882	1365	83158	4618682	55.1
3	0.0102	80761	823	80333	4453571	55.1	3	0.0105	82517	868	82065	4535524	55.0
4	0.0079	79938	631	79609	4373238	54.7	4	0.0083	81649	676	81297	4453458	54.5
5	0.0225	79307	1786	392067	4293629	54.1	5	0.0228	80973	1848	400245	4372161	54.0
10	0.0158	77520	1228	384531	3901562	50.3	10	0.0162	79125	1282	392420	3971916	50.2
15	0.0251	76292	1911	376682	3517031	46.1	15	0.0276	77843	2152	383834	3579496	46.0
20	0.0357	74381	2657	365261	3140349	42.2	20	0.0389	75691	2941	371101	3195662	42.2
25	0.0384	71724	2756	351729	2775089	38.7	25	0.0431	72750	3136	355908	2824561	38.8
30	0.0411	68968	2831	337761	2423360	35.1	30	0.0455	69614	3165	340155	2468652	35.5
35	0.0461	66137	3050	323058	2085599	31.5	35	0.0473	66449	3143	324385	2128497	32.0
40	0.0526	63087	3321	307129	1762541	27.9	40	0.0525	63306	3320	308227	1804112	28.5
45	0.0636	59765	3800	289325	1455412	24.4	45	0.0602	59985	3610	290900	1495886	24.9
50	0.0791	55965	4429	268753	1166086	20.8	50	0.0770	56375	4338	271030	1204985	21.4
55	0.1098	51536	5658	243536	897333	17.4	55	0.1048	52037	5452	246553	933955	17.9
60	0.1511	45878	6933	212057	653797	14.3	60	0.1411	46584	6575	216484	687402	14.8
65	0.2176	38945	8476	173534	441740	11.3	65	0.2003	40009	8013	180013	470917	11.8
70	0.3017	30469	9193	129361	268206	8.8	70	0.2866	31996	9169	137057	290904	9.1
75	0.4241	21276	9023	83821	138845	6.5	75	0.4011	22827	9156	91243	153848	6.7
80+	1.0000	12253	12253	55024	55024	4.5	80+	1.0000	13671	13671	62604	62604	4.6

(*Continued on next page*)

TABLE 8. New Life Tables for the White Population of the United States, 1780–1900 (*Continued*)

	White males, 1890–99							White females, 1890–99					
Age	q_x	l_x	d_x	L_x	T_x	e_x	Age	q_x	l_x	d_x	L_x	T_x	e_x
0	0.1305	100000	13052	91255	4948013	49.5	0	0.1106	100000	11059	92812	5121250	51.2
1	0.0316	86948	2750	85326	4856758	55.9	1	0.0292	88941	2599	87407	5028438	56.5
2	0.0142	84198	1199	83563	4771433	56.7	2	0.0140	86342	1209	85701	4941031	57.2
3	0.0092	82999	764	82602	4687870	56.5	3	0.0091	85132	775	84730	4855330	57.0
4	0.0071	82235	584	81931	4605268	56.0	4	0.0071	84358	603	84044	4770601	56.6
5	0.0202	81651	1649	404130	4523337	55.4	5	0.0196	83755	1646	414661	4686556	56.0
10	0.0139	80001	1109	397235	4119207	51.5	10	0.0136	82110	1117	407756	4271895	52.0
15	0.0219	78892	1731	390135	3721972	47.2	15	0.0231	80993	1873	400281	3864139	47.7
20	0.0318	77162	2454	379672	3331837	43.2	20	0.0328	79120	2594	389113	3463858	43.8
25	0.0347	74707	2590	367061	2952165	39.5	25	0.0369	76526	2825	375565	3074744	40.2
30	0.0378	72117	2725	353773	2585104	35.8	30	0.0396	73701	2920	361203	2699179	36.6
35	0.0429	69392	2975	339522	2231331	32.2	35	0.0418	70781	2957	346511	2337976	33.0
40	0.0492	66417	3266	323918	1891809	28.5	40	0.0467	67824	3165	331204	1991465	29.4
45	0.0597	63151	3768	306333	1567890	24.8	45	0.0542	64658	3505	314527	1660260	25.7
50	0.0748	59382	4439	285814	1261557	21.2	50	0.0700	61153	4281	295060	1345733	22.0
55	0.1043	54943	5729	260394	975744	17.8	55	0.0964	56871	5485	270646	1050673	18.5
60	0.1441	49214	7092	228342	715350	14.5	60	0.1311	51387	6739	240086	780027	15.2
65	0.2089	42123	8801	188610	487008	11.6	65	0.1889	44648	8435	202151	539941	12.1
70	0.2928	33321	9758	142213	298398	9.0	70	0.2741	36213	9926	156250	337790	9.3
75	0.4172	23564	9831	93242	156185	6.6	75	0.3903	26287	10260	105785	181540	6.9
80+	1.0000	13733	13733	62943	62943	4.6	80+	1.0000	16027	16027	75754	75754	4.7

TABLE 9. Proportion Dying in Age Interval x to $x + n$, White Males

Exact age, x	n	1860	1861	1862	1863	1864	1865	1866–69
0	1	0.2071	0.2071	0.2071	0.2071	0.2071	0.2071	0.2071
1	1	0.0469	0.0469	0.0469	0.0469	0.0469	0.0469	0.0469
2	1	0.0206	0.0206	0.0206	0.0206	0.0206	0.0206	0.0206
3	1	0.0132	0.0132	0.0132	0.0132	0.0132	0.0132	0.0132
4	1	0.0103	0.0103	0.0103	0.0103	0.0103	0.0103	0.0103
5	5	0.0295	0.0295	0.0295	0.0295	0.0295	0.0295	0.0295
10	5	0.0216	0.0216	0.0216	0.0216	0.0216	0.0216	0.0216
15	5	0.0339	0.0434	0.1300	0.1507	0.1682	0.1451	0.0339
20	5	0.0468	0.0549	0.1398	0.1821	0.2518	0.1703	0.0468
25	5	0.0491	0.0587	0.1440	0.1713	0.2230	0.1445	0.0491
30	5	0.0505	0.0588	0.1383	0.1755	0.2391	0.1617	0.0505
35	5	0.0555	0.0605	0.1118	0.1403	0.1927	0.1411	0.0555
40	5	0.0626	0.0656	0.0900	0.1069	0.1410	0.1148	0.0626
45	5	0.0747	0.0762	0.0829	0.0892	0.1032	0.0957	0.0747
50	5	0.0915	0.0955	0.0962	0.0971	0.0992	0.0990	0.0915
55	5	0.1251	0.1251	0.1251	0.1251	0.1251	0.1251	0.1251
60	5	0.1697	0.1697	0.1697	0.1697	0.1697	0.1697	0.1697
65	5	0.2394	0.2394	0.2394	0.2394	0.2394	0.2394	0.2394
70	5	0.3229	0.3229	0.3229	0.3229	0.3229	0.3229	0.3229
75	5	0.4402	0.4402	0.4402	0.4402	0.4402	0.4402	0.4402
80+		1.0000	1.0000	1.0000	1.0000	1.0000	1.0000	1.0000

from table 8 with a closely corresponding Haines life table. Figure 7 compares the 1870–79 life table for white males ($e_0 = 44.0$) with Haines's 1870 life table ($e_0 = 44.1$). In general, there is close correspondence between the two age profiles. The 1870–79 life table indicates higher mortality rates between ages 10 and 30, but the difference is modest: age-specific mortality rates are 12 percent higher at age 20 than in Haines's table. Figure 8 compares the 1840–49 life

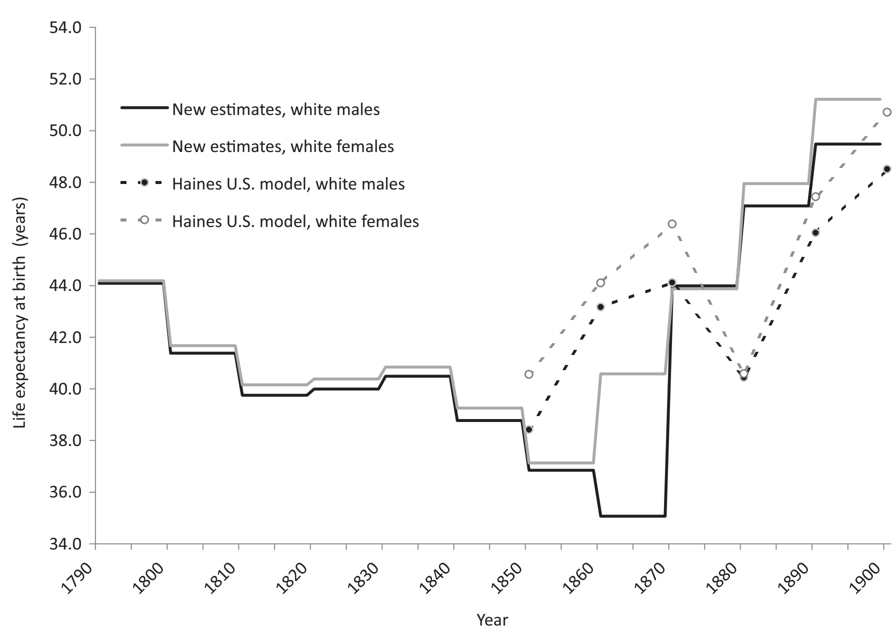

FIGURE 6. Life expectancy at birth, white population of the United States, 1790–1900.

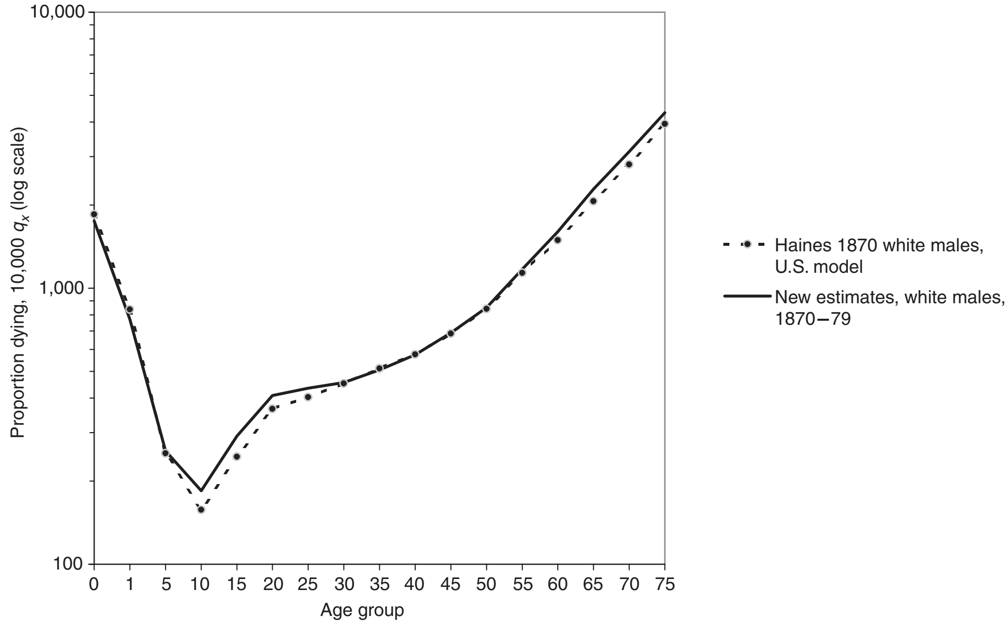

FIGURE 7. Proportion dying by age group, white males.

table for white females ($e_0 = 40.6$) with Haines's 1850 life table ($e_0 = 40.6$). Again, with the exception of the part of the profile between adolescence and middle age, there is close correspondence between the two age profiles. The difference between the curves between ages 10 and 35, however, is much greater. At age 20 the 1840–49 life table suggests a mortality rate 27 percent higher than Haines's 1850 life table.

Figure 8 also includes a modified plot of the proportions dying in Haines's 1850 female life table. The age-specific

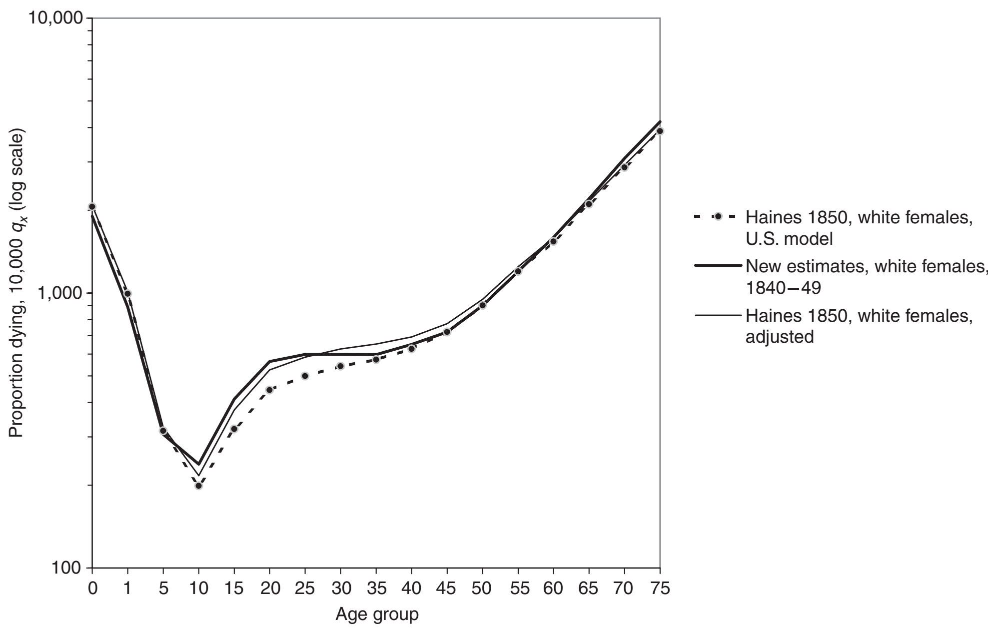

FIGURE 8. Proportion dying by age group, white females. q_x = proportion dying in age interval.

death rates in the Haines table were modified by doubling cause-specific death rates attributed to pulmonary tuberculosis and maternal mortality reported by Samuel Preston for national populations with life expectancy at birth less than 45 years (1976). The adjusted profile corresponds very closely with the age profile of female mortality in the 1840–49 life table. Although speculative and seemingly large, the adjustments correspond with what we know about changes in mortality and fertility between the mid-nineteenth and early twentieth centuries. As discussed earlier, mortality from tuberculosis fell rapidly in the late nineteenth century. Fertility among women in the 1900–1902 DRA life table was approximately half that of women in the 1840s. Although we cannot know the true age-specific mortality rates for white women in the 1840s, it is likely that the profile differed from that in the 1900–1902 DRA in the way indicated.

All of the decennial tables in table 8 are based, of course, on assumptions with substantial risk of error. Much more research is needed on biases in demographic estimation from genealogical sources. Although based in part on comparison with other sources and in part on the suspected impact of migration censoring and selection biases, the crude assumptions about the overestimation of adult life expectancy in genealogical-based estimates and sex differentials in adult life expectancy made in tables 4 and 5 are sources of potential error. Another weakness is the required method of inferring a complete life table from a single parameter, life expectancy at age 20. Historical demographers in Europe and elsewhere have called attention to the changing relationship between infant, childhood, and adult mortality over the course of the nineteenth century (Woods 1993). It is unlikely that the United States was an exception. Empirical research on infant and childhood mortality in the United States is sorely needed. Source material, however, remains a major issue.

Despite these caveats, the life tables shown in table 8 should prove useful for a wide variety of historical research. In addition to capturing known mortality trends not reflected in existing life tables, they more accurately represent the likely sex- and age-specific profile of nineteenth-century mortality. The life tables should also prove useful as a point of reference for subsequent studies and critiques. With any luck, nineteenth-century demographers will have more choices of life tables with a firmer empirical base in the not too distant future.

NOTES

1. This work was supported in part by NIHCD grant number 1 K01-HD052617–01 and an Arthur H. Cole Grant-in-Aid Award from the Economic History Association. The author would like to thank Samuel H. Preston, Douglas Ewbank, and Michael R. Haines for helpful comments.

2. The original Death Registration Area (DRA) included only 10 states and the District of Columbia. The system was deemed complete in 1933, when Texas was added to the system, although considerable underreporting of births and deaths continued to plague the system until the 1940s.

3. Although Massachusetts's death registration system was implemented in 1842, it took several years for the system to become effective. By 1860, Robert Gutman has estimated that only 8 percent of deaths were unrecorded (Vinovskis 1972, 186).

4. Condran and Crimmins's (1980, 188–90) application of the Chandra Sekar-Deming technique suggests that approximately 85 percent of deaths in rural areas and 92 percent of deaths in urban areas were registered. Infant deaths were missed more often than deaths at other ages.

5. Other potential problems include the possibility that the deaths of children ages 5–19, while more fully enumerated than deaths at other ages, were still underreported, and the possibility of a changing level of undercount from census to census. If underreporting was significant, the Haines (1998) life tables may overstate life expectancy. The addition of some state death registers in 1880 likely lowered the overall undercount and may explain some of the sharp decline in life expectancy between the 1870 and 1880 estimates.

6. Kasakoff and Adams (1995) report the average age at death by birth cohort, not period. In the figure, the cohort estimates are offset 20 years to increase comparability.

7. The urban population is defined liberally as all individuals living in urbanized areas and in all places of 2,500 or more residents outside of urbanized areas. The percentage living in large cities with significant sanitation problems was much smaller. The urban population increased from 6.1 percent in 1800 to 10.8 percent in 1840, 28.2 percent in 1880, and 51.2 percent in 1920. By the turn of the century, when urbanization was significant enough to pose a major impact on national life expectancy, the public health movement had made significant strides in introducing clean water supplies, sewer systems, and other public health projects, greatly reducing the urban-rural differential in life expectancy.

8. Considerable uncertainties surround estimates of real national income in the early nineteenth century. Most economic historians conclude that there was a sharp increase in real economic growth in the 1820s. According to Richard Sutch, the annual growth rate between 1800 and 1828 averaged about 0.6 percent per year. Between 1828 and 1860 it averaged more than twice that rate (Sutch 2006).

9. Under some conditions, censoring bias does not impart a downward bias in life expectancy estimates. If a researcher knows when an individual disappeared from observation and if censored individuals experienced the same risk of death as noncensored individuals, for example, it is possible to construct nonbiased age-specific mortality estimates. Relative to the extensive rules followed by analysts of community-based reconstitution studies, however, researchers relying on genealogical data have shown little interest in precisely determining when the population was under observation. Neither Kunze (1979) nor Pope (1992) appeared to have included risk years from right-censored individuals in the calculation of age-specific death rates. Only individuals with known birth and death dates are included. Given these selection criteria, censoring bias will impart a downward bias (for an extended rumination on biases in early American mortality studies, see Smith 1979).

10. The average of Haines's (1998) 1850 and 1860 U.S. model census-based estimates of life expectancy at age 20 was assumed to be representative of the 1850s, the 1870 and 1880 estimates representative of the 1870s, and the 1880 and 1890 estimates representative of the 1880s. It was not assumed that the average of the 1860 and 1870 estimates would be representative of the 1860s, however, because the census-based estimates fail to consider the impact of the U.S. Civil War (1861–65).

11. Although genealogies are successful in tracking some family members across time and space, migrating family members are more prone to be lost from observation. Hall and Ruggles (2004) have shown that internal migration in the United States exhibited a "U-shaped" pattern between 1850 and 2000. Almost one-in-two whites age 50–59 between 1850 and 1880 were living in a state other than their birth state. This ratio dropped steadily after 1880, reached a low of about one-in-three in the period 1940–70 and then increased to over four-in-ten in the 2000 census.

12. Although Kunze's (1979, 200) sample appears to be slightly larger than Pope's (1992, 282) sample, Kunze does not report the number of cases used in his period estimates. The combined estimates shown in table 5 are therefore unweighted averages, smoothed slightly in the period before 1850.

13. Estimates of the white birth rate were obtained with stable population methods, the published age distributions of the 1800 census, and life tables constructed by fitting the adjusted and unadjusted Pope (1992) and Kunze (1979) estimate of life expectancy at age 20 to the 1901 rural DRA life table as described in the latter part of this article.

14. Males and females enter Pope's (1992) sample as either a child of bloodline parents or as a spouse of a bloodline individual. The former

contribute risk years from birth to death while the latter contribute risk years from marriage to death. Theoretically, there should be approximately equal numbers of men and women in the samples. According to Pope's illustration of a "typical" family history, however, 11 percent of men in the genealogical samples had a missing birth date and 43 percent a missing death date. The percentages for women were 16 and 59 percent, respectively (ibid., 273). As a result, Pope's period life expectancy estimates are based on 3,166 males and 2,338 females with known birth and death dates (ibid., 282). Kunze (1979, 200–204) does not discuss the completeness of his demographic data by sex, but similar differences are apparent in the number of males and females used in his analysis. Kasakoff and Adams (1995) report only the mean age at death of males.

15. Stolnitz's (1956, 23–25) classic review of long-term mortality trends called explicit attention to instances of higher female mortality in the previous century. Although females in Western countries between 1840 and 1910 typically enjoyed lower mortality rates during infancy and older ages, higher female mortality rates from late childhood through most of the childbearing years was common. The modern pattern of lower female mortality at all ages did not become typical until the 1930s. Although the life tables Stolnitz examined tended to favor higher female life expectancy at all ages, higher male life expectancy could be found across an "appreciable" range of ages well into the twentieth century in Ireland, Italy, Austria, and Bulgaria.

16. Stolnitz (1956, 23–25) reported the largest persisting female disadvantages in life expectancy among the Irish population, which experienced high fertility, low nutritional status, preferential treatment for males, and endemic tuberculosis well into the twentieth century (ibid.; Kennedy 1973).

17. The results also suggest that Haines's (1998) life tables overstate female life expectancy at age 20 relative to male life expectancy. The relative overstatement is likely a result of Haines's choice of model life tables—a "U.S. model" constructed from the 1900–1902 DRA and Coale, Demeny, and Vaughan's (1966) west model. Both models are based on the mortality experience of more urban and lower fertility populations than the nineteenth-century population of the United States. As discussed at greater length in the section on the age profile of nineteenth-century mortality, these models likely understate female mortality during childbearing years relative to other ages and overstate female life expectancy at age 20 relative to male life expectancy.

18. Details on weighting and combining the 1900–1902 overall and rural DRA life tables can be found in the section on new decennial life tables and in note 24 below.

19. Four- and five-parameter models have also been proposed (see, e.g., Ewbank, de Leon, and Stoto 1983).

20. Although Coale and Zelnik (1963, 168–69) observed a good correspondence between the 138 life tables that were used to construct the west model and the 1900–1902 DRA life table, only 36 of the 138 life tables came from nineteenth-century populations. The model matches the male experience better than the female experience. Coale and Zelnik did not compare the 1900–1902 rural DRA life table with the model.

21. Rebecca Kippen (2005) has noted that maternal deaths are often underreported in official statistics and in estimates derived from family reconstitution studies. Her revised estimates of maternal mortality for nineteenth-century Tasmania—7 deaths per 1,000 live births—are approximately twice as high as estimates derived from other sources. Even so, maternal mortality remained a distant second leading cause of death among women age 29–44 behind pulmonary tuberculosis.

22. It is important to remember that the 1900–1902 overall DRA included females in the rural DRA. The differences would have been greater if we were able to compare urban females directly to rural females (for an analysis of urban-rural mortality differentials in 1890 and 1900 see Condran and Crimmins 1980).

23. Typically, sex mortality differentials favor females at lower mortality levels. Sex differences in mortality between historical and modern populations are the result of changes in causes of death associated with mortality decline. Female advantages in mortality at all ages emerged only with the decline of tuberculosis and other infectious diseases as leading causes of death and their replacement with degenerative diseases. The decline of maternal mortality also played a small role (Preston 1976).

24. According to table 2, the 1900–1902 DRA life table was 60.1 percent urban and the 1900–1902 rural DRA life table was 13.2 percent urban. If W1 is the weight needed for the overall DRA life table, W2 is the weight for the rural DRA life table, and the desired combined life table is 40.2 percent urban, then $(W1 \times 60.1) + (W2 \times 13.2) = 40.2$. Further, $W1 + W2 = 1$. Solving the second equation for W2, we get W2 = 1—W1. By substitution, the first equation becomes $(W1 \times 60.1) + ([1- W1] \times 13.2) = 40.2$. Solving for W1, we get 0.575. Substituting the result in the second equation and solving yields 0.425 for W2.

25. Haines and Preston (1997, 77) state that the "improvement was most rapid in large urban areas, where mortality had been the worst. The substantial urban mortality penalty . . . of the late nineteenth century was rapidly disappearing by the early twentieth century. Public health improvements, better nutrition and shelter, and some advances in medical science all played a role."

26. Cohort differences between the male-female *differential* in 10-year survivorship ratios in the 1860s relative to the average male-female differential in 10-year survivorship ratios in the 1850s and 1870s were assumed to be because of the excess male mortality in the war. The estimate required four major assumptions: (1) the native-born white population was closed to migration; (2) changes in net census underenumeration had an equal impact on native-born white males and females; (3) foreign-born white men suffered rates of mortality in the war equal to the native-born white population; and (4) there were negligible civilian deaths among native-white women age 15–45. For the approximately equal rates of mortality among foreign-born and native-born men, see Lee (2003, 60). For the limited number of civilian casualties in the U.S. Civil War, see McPherson (1988, 619) and Neely (2007). Although the resulting estimate of approximately 713,000, excess male deaths is larger than the 588,000 usually attributed to white men in the war, there are many reasons to assume the 588,000 figure is too low (Hacker 1999, chap. 2; Faust 2006).

27. The Union Army data set, collected by the University of Chicago Center for Population Economics and Brigham Young University under the direction of Robert W. Fogel, was used to parse deaths by year and within five-year age groups.

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