

# The Rise in Old-Age Longevity and the Market for Long-Term Care

By DARIUS LAKDAWALLA AND TOMAS PHILIPSON\*

In many countries around the world, the demographic transition into reduced fertility and mortality has forced the private and public sectors to grapple with the care of rapidly aging populations. Since 1960, the share of the U.S. population above 65 years of age has grown substantially, from about 9 percent to 14 percent. Other developed countries have experienced even more rapid growth. For example, in many European nations, the elderly population accounts for nearly one-fifth of the total population, and growth in this share has been larger than in the United States over the past few decades. As the elderly population has grown, the share of GDP devoted to long-term care for the elderly has grown as well. This rapid growth has stimulated interest in the study of how private markets for long-term care function and how they are affected by various forms of public intervention.

Concern about the importance of long-term care has further intensified, because several key market forces in the United States have combined to exert tremendous upward pressure on the market output of long-term care. First, the

share of output that is publicly financed by Medicaid has grown enormously, from about 24 percent of 1971 nursing home bed-days to about 65 percent of 1991 bed-days.<sup>1</sup> Second, from 1920 to 1940, birth rates fell by almost 30 percent.<sup>2</sup> Compared to her 1970 counterpart, a 70-year-old in 1990 had far fewer children to care for her in lieu of a nursing home. Third, compounding the effects of declining fertility, the rising human capital, wages, and labor-force participation of young to middle-aged women has increased the opportunity cost of family-produced care, which is provided predominantly by daughters rather than sons. In 1970, 53 percent of women between the ages of 40 and 55 participated in the labor force, but by 1990 this labor-force participation rate had climbed to 74 percent.<sup>3</sup> Finally, towards the mid-1980's, entry and investment barriers in the nursing home industry, erected by Certificate of Need laws, were relaxed considerably; the resulting increase in the supply of nursing home beds served to expand the output of long-term care even more.<sup>4</sup> Since this combination of forces has worked to drive up per capita market demand and supply, one would suspect that long-term care output should have grown faster than the elderly population.

However, Figure 1 demonstrates that this has not been the case in the United States over the last few decades.<sup>5</sup> The figure compares the

\* Lakdawalla: RAND, 1700 Main Street, Santa Monica, CA 90407, and National Bureau of Economic Research (e-mail: [darius\\_lakdawalla@rand.org](mailto:darius_lakdawalla@rand.org)); Philipson: Harris School of Public Policy, Department of Economics, and The Law School, University of Chicago, 1155 E. 60th Street, Chicago, IL 60637 (e-mail: [t-philipson@uchicago.edu](mailto:t-philipson@uchicago.edu)). We are thankful for the constructive comments made by Gary Becker, Thomas DeLeire, Will Dow, Martin Gaynor, Paul Gertler, Chulhee Lee, Casey Mulligan, Edward Norton, Richard Posner, James Vaupel, and an anonymous referee. Helpful input was also provided by seminar participants at the 1998 AEA meetings, the 1998 Midwestern Economic Association meetings, the 1998 NBER Summer Institute, the University of California-Berkeley, the University of Chicago, the American Enterprise Institute, and the Max Planck Institute, Germany. Exceptional research assistance was provided by Hyo Joon You, David Grabowski, and Anup Malani. Philipson gratefully acknowledges financial support from National Institute on Aging (Grant No. R01-AG16494) and the Research Fellows Program of the Alfred P. Sloan Foundation.

<sup>1</sup> See National Center for Social Statistics (NCSS) (1974) for the 1971 Medicaid data. 1991 data are based on HCIA (1996). For a discussion of why Medicaid subsidization has grown so much over this time period, see Philipson and Lakdawalla (2001).

<sup>2</sup> See U.S. Department of Commerce, Bureau of the Census (1965), series B19-30 and B31-36.

<sup>3</sup> Labor-force participation rates are constructed from the Public Use Microdata Samples of the 1970 and 1990 U.S. Census. See Steven Ruggles and Matthew Sobek (1997) for the raw data.

<sup>4</sup> See Charlene Harrington et al. (1997).

<sup>5</sup> The figure uses data on nursing home residents from the 1973, 1977, 1985, 1991, and 1995 National Nursing

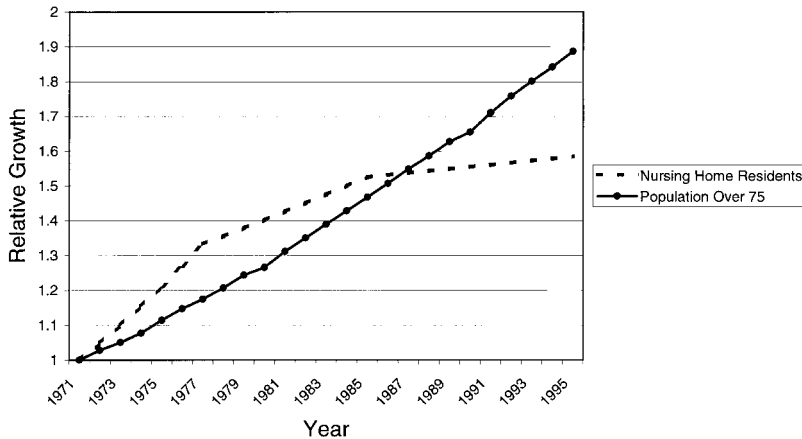


FIGURE 1. RELATIVE GROWTH OF NATIONWIDE NURSING HOME RESIDENTS VERSUS RELATIVE GROWTH OF ELDERLY POPULATION

growth since 1971 in current nursing home residents to the growth in the population over the age of 75.<sup>6</sup> From Figure 1, we learn that growth in nursing home residents has rapidly decelerated since 1971 in spite of constant rates of elderly population growth: specifically, in the mid-1970's, the resident population grew at a 4.8-percent annual rate; by the early 1980's, this annual growth rate had plummeted by almost two-thirds to 1.7 percent; finally, during the late 1980's and early 1990's, the growth rate dropped further, to a mere 0.4 percent per year. Nursing home growth has decelerated in spite of relatively stable growth rates for the elderly population: the population over 75 has grown at a roughly constant annual rate of 2.7 percent for the past two decades. As a result of these divergent trends, the resident population grew at twice the rate of elderly population during the 1970's, but at less than half of this rate during

the 1980's and early 1990's. In fact, per capita output contracted so sharply during the 1980's that it more than offset the growth that occurred during the 1970's. It is remarkable that, from 1971–1995, per capita output fell by almost 20 percent overall, in spite of the many forces which should have pushed it upwards. Figure 1 thus poses the central question of this paper: why did per capita output change course during the 1980's, and how did it contract so sharply in the presence of many forces that should have pushed it higher?

Motivated by this question, this paper analyzes the relationship between aging and the growth of long-term care markets. We argue that aging may actually *decrease* the per capita demand for market care if it raises the supply of nonmarket care produced by other elderly persons. This effect may be counteracted or reinforced by changes in the incidence of disability among the elderly. However, the important point is that disability reduction has not only a direct negative effect on nursing home demand, but also an *indirect* supply effect, because it expands the supply of nonmarket care by other elderly people. Significantly, an elderly person who has no demand for care may in fact become a supplier of care. These indirect supply effects have the surprising implication that market care may *contract* with the longevity of the scarcer sex, typically males, and expand with the *healthy* life expectancies of the abundant sex, typically females. Since men tend to die before

Home Surveys. (The raw data are reported in Genevieve W. Strahan, 1997.) Intermediate years are interpolated, and data for 1971–1972 are extrapolated, assuming constant rates of growth between observed points. Population data from 1970 to 1995 come from the National Center for Health Statistics website: <http://www.cdc.gov/nchswww/datawh/statab/unpubd/mortabs/pop7095.htm>. The 1971 baseline values are: 977,481 nursing home residents, and 7,877,000 people over age 75.

<sup>6</sup> Most consumers in the long-term care market are above the age of 75. In 1995, about 17 percent of residents were 65–74, 42 percent between 75–85, and 41 percent above 85 years (National Center for Health Statistics [NCHS], 1995).

women, growth in elderly males causes couples to stay married longer, raises the supply of spousal care, and thus lowers the demand for market care. Conversely, relative growth in healthy elderly females causes more women to spend time in widowhood, where they have no spouse to care for them, and thus raises the per capita demand for market care.

We examine these implications empirically using a panel data set containing all U.S. counties over the last three decades. We find evidence consistent with our prediction that growth in elderly men will reduce market-based long-term care: a ten-percentage-point increase in the ratio of men per woman appears to reduce the per capita stock of nursing home residents by as much as 16 percent.<sup>7</sup> As for Figure 1, we argue that relative declines in the number of elderly men contributed to per capita output growth during the 1970's, while the decline in demand during the 1980's and 1990's resulted from improvements in the health of the elderly. Changes in sex composition played almost no role over this period, as elderly men and women grew at roughly the same rate. We estimate that 60 percent of the per capita growth during the 1970's was caused by relative growth in elderly females, while 70 percent of the per capita decline during the 1980's was caused by improvements in the health of the elderly. The next most significant force, the dramatic expansion in Medicaid subsidization, explained only 15 percent of the per capita growth during the 1970's and played no role during the 1980's.

This paper relates to an existing body of work on long-term care, but relatively little analytic attention has been paid by economists to the macroeconomic aspects of the market for long-term care in general, or the aggregate impact of aging on this market in particular. For a review of the existing long-term care literature, see

Edward C. Norton (1999). For an overview of policy issues and this market, see, for example, Alan M. Garber (1994) or Norton and Joseph P. Newhouse (1994). Analyses of the empirical determinants of nursing home entrance by individuals are provided, for example, by Garber and Thomas MaCurdy (1990), Axel H. Börsch-Supan et al. (1991), and Steven N. Stern (1995).

### I. The Demographics and Economics of Long-Term Care

The output of long-term care is driven ultimately by the presence of disabled people whose best option for care is a nursing home. To study the determinants of disability, let  $T$  represent total lifetime in years, with life expectancy given by  $\mu \equiv E[T]$ ;  $S$  is the number of years an individual spends in disability, with expected value  $\nu \equiv E[S]$ ; consequently,  $T - S$  represents healthy years, while  $\mu - \nu$  represents expected healthy life span. The two durations have the survival curves depicted in Figure 2.<sup>8</sup> The top survival function is for the overall lifetime survival  $S_T(t)$ , the probability that an entrant is alive at time  $t$ , while the bottom survival is the healthy lifetime survival  $S_{T-S}(t)$ , or the probability that an entrant is alive and healthy at time  $t$ .

The frail members of a cohort have completed their healthy durations but not their life. If the cohort size is  $\lambda$  then the frail stock  $F$ , the black area between the two curves in Figure 2, satisfies:

$$(1) \quad F \equiv \lambda \int [S_T(t) - S_{T-S}(t)] dt = \lambda \nu.$$

In other words, the number of frail people is equal to the size of each cohort, multiplied by the number of years each cohort is expected to spend in frailty. Similarly, the gray area below the lower curve in Figure 2 represents

<sup>7</sup> In Lakdawalla and Philipson (1999), we also investigated whether or not the county-level evidence on the sex ratio effect is consistent with individual-level evidence. Using microdata, we found that the presence of a spouse *more than halves* the probability of nursing home entrance. Furthermore, not only the signs, but also the magnitudes of the individual-level effects of marriage on market care were found to be consistent with the county-level estimates.

<sup>8</sup> Using an assumption of a constant hazard rate into death, the  $S_T(t)$  curve is constructed from expected longevity in 1991, which is 11.1 for 75-year-olds (U.S. Department of Health and Human Services, *Vital Statistics of the United States*, 1991). The  $S_{T-S}(t)$  curve is the disability incidence at age  $t$  multiplied by the probability  $S_T(t)$ . Disability incidence is constructed using 1994 data for 65–74-year-olds, 75–84-year-olds, and those over 85 (Kenneth G. Manton et al., 1997).

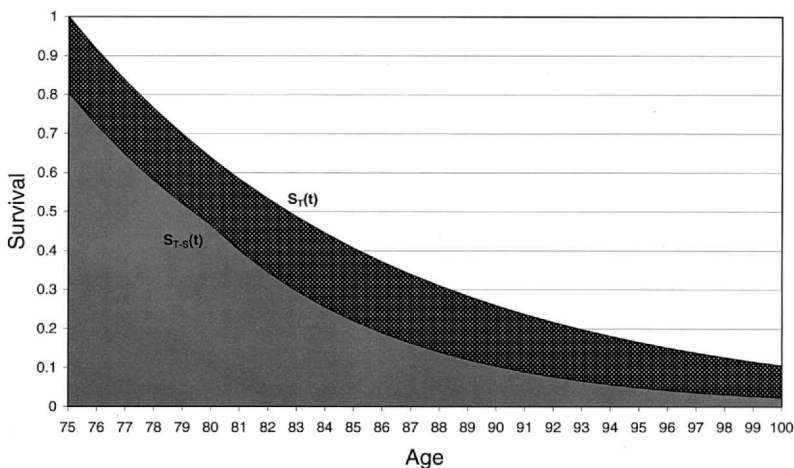


FIGURE 2. OVERALL AND HEALTHY LIFETIME SURVIVAL FUNCTIONS

the stock  $H$  of healthy individuals, which satisfies:

$$(2) \quad H \equiv \lambda \int S_{T-s}(t) dt = \lambda(\mu - \nu).$$

We consider long-term care, whether produced in the household or in the market, as the continuing care of an individual who has reached frailty. The total demand  $D$  for market care is the stock of frail multiplied by their per capita demand,

$$(3) \quad D(p, H, F) \equiv Fd\left(p, \frac{H}{F}\right).$$

The function  $d\left(p, \frac{H}{F}\right)$  denotes the per capita nursing home demand and is assumed to fall in the price of nursing home care and the availability of healthy nonmarket caregivers:  $d_1 \leq 0$  and  $d_2 \leq 0$ . We assume that the aggregate supply of market-based care slopes upward and abstract from technological change, so that the supply function remains fixed over time. We focus on movements in demand as the key source for observed changes in output.

We may express the aggregate demand for

market care in terms of those population parameters alone, as in

$$(4) \quad D(p, \lambda, \mu, \nu) = \lambda \nu d\left(p, \frac{\mu}{\nu} - 1\right).$$

This allows us to decompose growth in demand, at a given price, into changes driven by each of these underlying parameters:

$$(5) \quad \left. \frac{dD}{D} \right|_p = \frac{d\lambda}{\lambda} \varepsilon_\lambda + \frac{d\mu}{\mu} \varepsilon_\mu + \frac{d\nu}{\nu} \varepsilon_\nu,$$

where, generically,  $\varepsilon_x \equiv \frac{\partial D}{\partial x} \frac{x}{D}$  denotes the elasticity of aggregate market demand with respect to the variable  $x$ . This decomposes demand growth into an entry component, a longevity component, and a disability component. According to equation (4), the elasticity of entry is unity,  $\varepsilon_\lambda = 1$ , because a rise in entry simply expands the healthy and frail populations proportionally. The elasticity with respect to longevity is *negative*,  $\varepsilon_\mu \leq 0$ , as it raises the stock of healthy people who can care for the frail, but not the stock of frail people. This is the central negative effect of aging on market care stressed in this paper. In terms of Figure 2, growth in longevity alone will expand the light

area while leaving the dark area unchanged. The elasticity of expected disability is positive:  $\epsilon_\nu \geq 0$ , as an expansion in disability simultaneously raises the stock of frail people *and* makes suppliers of home care more scarce. If disabled life span grows more slowly (quickly) than overall life span, long-term care demand will grow more slowly (quickly) than disabled life span.

When men and women are matched together, the population parameters determine not only the stock of frail males and females, but also their marriage patterns. Since men are the scarcer sex, relative growth in elderly males raises the prevalence of marriage among the elderly. This provides greater access to spousal care and thus reduces the demand for market-based long-term care. Based on these ideas, in Lakdawalla and Philipson (1999) we adapt equation (4) to include the effect of marriage. Denote the entry and longevity of elderly males as  $\lambda^m$  and  $\mu^m$ , and let  $\lambda^f$  and  $\mu^f$  denote the analogous variables for elderly females. This allows us to write:

$$(6) \quad D\left(p, \lambda, \mu, \nu, \frac{\lambda^m \mu^m}{\lambda^f \mu^f}\right) \\ = \lambda \nu d\left(p, \frac{\mu}{\nu} - 1, \frac{\lambda^m \mu^m}{\lambda^f \mu^f}\right),$$

where  $d_3 < 0$ , because relative increases in elderly males lower per capita demand.

### II. Empirical Analysis

The empirical trends in long-term care shown in Figure 1 can be explained by two forces: the relative growth of elderly females during the 1970's, and improvements in elderly health during the 1980's. In the early 1970's, the male population over age 75 grew at a mere 1.7-percent annual rate, while the female population grew at a 3.4-percent annual rate; we will argue that this disparity accounted for the dramatic rise in per capita market care output witnessed during the 1970's. By the early 1980's, this huge imbalance had been largely wiped out, as men were then growing at a 2.6-percent annual rate, while women were growing at just a 2.9-percent rate. As the 1980's progressed, the male growth rate caught up to and eventually even

surpassed the female growth rate. During the 1970's, the ratio of males to females, which roughly represents the share of women married, fell from 0.64 in 1970 to 0.55 in 1980 for the over-75 age-group.<sup>9</sup> As a result of this decline, there were about 900,000 more unmarried elderly women in 1980 than there would have been at the 1970 rate of marriage. This increase in widowhood, substantial in relation to the 1.4 million nursing home residents in 1980, helped push up the per capita output of market care during the 1970's. While changes in the sex ratio became less significant during the 1980's, improvements in elderly health became more significant during this decade. In 1981, the incidence of disability among the population over 75 was 31.9 percent, while in 1991, this rate fell to 28.1 percent.<sup>10</sup> Since there were about 13.5 million people over age 75 in 1991, there were roughly half a million fewer disabled persons over 75 in 1991 than there would have been absent the disability reduction. This represents a highly significant share of the 1.5 million residents of nursing homes in 1991.<sup>11</sup>

To test the restrictions placed on equation (6), we use data on the output of nursing home residents and the county-level populations of the elderly. Our predictions for demand growth can be translated into predictions for output data, because output growth is aggregate demand growth multiplied by the price effect:

$$\frac{dY}{Y} = (1 + \epsilon_p) \frac{dD}{D} \Big|_p, \text{ where } \epsilon_p \equiv$$

$$\frac{\frac{\partial D}{\partial p}}{\frac{\partial Z}{\partial p} - \frac{\partial D}{\partial p}}.$$

Unfortunately, population data do not allow us to separate the output effects of changes in total longevity,  $\mu$ , from the effects of frail life span,  $\nu$ , but we can separate the effects

<sup>9</sup> All calculations were performed using male and female age-specific population data from the 1970–1991 issues of the U.S. Department of Health and Human Services publication, *Vital Statistics of the United States*.

<sup>10</sup> These rates are calculated from disability incidence estimates for people of ages 75–84, and over 85; all these estimates are taken from Manton et al. (1997). These are converted into rates for people over 75 by using population data from *Vital Statistics of the United States*.

<sup>11</sup> The data on nationwide nursing home residents are taken from Strahan (1997).

of longevity from those of changes in the incidence of disability,  $\frac{\nu}{\mu}$ , which equals the steady-state proportion of the elderly population that is disabled. Consider the panel regressions of the following form, where the unit of observation is county  $i$  in year  $t$ :<sup>12</sup>

$$(7) \quad \ln(Y_{it}) = \beta_0 + \beta_1 \ln(\lambda_{it}\mu_{it}) \\ + \beta_2 \left( \frac{\lambda_{it}^m \mu_{it}^m}{\lambda_{it}^f \mu_{it}^f} \right) + \beta_3 \mathbf{X}_{it} + \phi_i + \eta_t + \varepsilon_{it}.$$

Assuming a steady-state in elderly males and females,  $\lambda_{it}\mu_{it}$  can be measured as the elderly (over 65 or 75) population of county  $i$  at time  $t$ ,<sup>13</sup> and  $\frac{\lambda_{it}^m \mu_{it}^m}{\lambda_{it}^f \mu_{it}^f}$  can be measured as the ratio of males to females in county  $i$  at time  $t$ .  $\mathbf{X}_{it}$  represents a vector of variables characterizing government regulation of the nursing home industry in county  $i$  at time  $t$ . The regression includes a county-specific fixed effect  $\phi_i$ , and a year-specific fixed effect  $\eta_t$ . We assume that the incidence of disability changes over time, but not across counties within the same year.<sup>14</sup> Therefore, the effect of disability changes on demand is absorbed by  $\eta_t$ , the year fixed ef-

fect.<sup>15</sup> Given these assumptions, the demand function in equation (6) allows us to interpret the coefficients.  $\beta_1$  measures the elasticity of output with respect to population,  $\lambda\mu$ , holding the sex ratio and the incidence of disability fixed. According to equation (6), demand grows proportionally with  $\lambda$  and  $\mu$ , at a fixed disability incidence and sex ratio.<sup>16</sup> This implies that  $\beta_1 = 1 * (1 + \varepsilon_p)$ , and that  $0 \leq \beta_1 < 1$ . Equation (6) also implies that  $\beta_2$ , the partial elasticity with respect to the sex ratio, is negative, because increases in the sex ratio lower demand.

There are two ways to estimate equation (7). We report results for both methods. First, we run the regression within counties and years. Second, we first-difference all the data, to remove the county fixed effect, and then add year dummies to remove the remaining year-specific effects. The data we use for this analysis come from the Area Resource File (U.S. Department of Health and Human Services, 1996). From this file, we have taken data on the long-term care output and demographic characteristics of every county in the United States. The file contains county-level data on the number of long-term care facilities, the total number of residents in all such facilities, and the number of men and women over the ages of 65 and 75. These data are available for 1971, 1973, 1976, 1978, 1980, 1982, 1986, and 1991. Unfortunately, the definition of long-term care facilities changes slightly over this period. For 1971–1978, a long-term care facility is defined as a nursing home or a personal care home, and the data come from the National Master Facility Inventory (NMFI). For 1980–1982, the data include only nursing homes and also come from the NMFI. For 1986–1991, the data include nursing and board/care (otherwise known as residential care) homes. These differences appear not to be significant, because our results do not seem sensitive to the years

<sup>12</sup> We do not include a variable for price, because county-level price data are not available. Even state-level price data cannot be obtained: while the National Nursing Home Surveys report price at the nursing home level, they do not release the state identifiers that would be necessary to link price and demographic data together.

<sup>13</sup> Another potentially interesting empirical specification would split apart the elderly population into its respective age categories and estimate the effects of altering the entire age distribution rather than just its mean. For example, instead of just  $\lambda_{it}\mu_{it}$ , we would have several variables, one for population between 65 and 70, one for population between 70 and 75, and so on. Of course, it would require much further theoretical work which we do not attempt here to pin down the theoretical sign of these coefficients, which reflect the elasticities of increases in one age-group, holding constant the sizes of all other age-groups.

<sup>14</sup> There exists some evidence that disability has not grown as fast as longevity across cohorts. For example, Manton et al. (1997) find that, holding the age distribution constant, the percentage of healthy persons (without measured disabilities) over the age of 85 has risen from 34.8 percent in 1982 to 40.2 percent in 1994. Also see Eileen M. Crimmins et al. (1997).

<sup>15</sup> Unfortunately, we cannot explicitly measure disability at the county level, due to data constraints. However, in Lakdawalla and Philipson (1999) we show how individual-level disability data may be used to identify empirically the effects of changes in health.

<sup>16</sup> Holding the incidence of disability fixed means that  $\nu$  grows at the same rate as  $\mu$ .

used. The population data on elderly males and females can be broken down into males and females over the age of 65 and over the age of 75. These data come from the 1970, 1980, and 1990 Census. To make them comparable with the long-term care data, the population data were exponentially interpolated by county to construct series for 1971, 1973, 1976, 1978, 1980, 1982, 1986, and 1991. The first government policy variable (in  $X_{it}$ ) at our disposal is the statewide share of bed-days subsidized by Medicaid. The 1991 values for this series are taken from HCIA (1996). The 1971 data are constructed by taking the total statewide number of Medicaid bed-days (National Center for Social Statistics, 1974) and dividing this by an estimate of total statewide 1971 bed-days from the Area Resource File. Intermediate years are then linearly interpolated. In addition to the Medicaid data, we have data for 1978, 1982, 1986, and 1991 on the presence of statewide Certificate of Need (CON) laws and statewide moratoria on nursing home bed construction from Harrington et al. (1997).<sup>17</sup> These represent additional components of the vector  $X_{it}$ . The data are summarized in Table 1.

As the table demonstrates, there was a significant decrease in the number of elderly men per woman over this 20-year period. We should note that nearly all of this decrease took place during the 1970's. While the table also shows a reduction in the incidence of CON laws, the table understates the extent of the reduction, because during the 1980's, CON laws became much less aggressively enforced.<sup>18</sup>

The results of the panel regressions are reported in Tables 2 and 3. The results of within-county and within-year estimation are reported in Table 2, while the results of the first-differencing method are reported in Table 3. Regardless of how we measure the elderly population, either as the population over 65 or 75, and regardless of which government policy we include, our predictions that  $0 \leq \beta_1 \leq 1$  and  $\beta_2$

TABLE 1—MEAN ATTRIBUTES OF LONG-TERM CARE IN U.S. COUNTIES (1971–1991)

	1971	1991
Nursing home residents	349	582
	1,194	1,605
Population over 65	6,699	10,322
	22,757	31,023
Men per woman over 65	0.81	0.71
	0.23	0.09
Female marriage rate over 65 <sup>a</sup>	NA	0.41
		0.06
Population over 75	2,561	4,365
	8,566	13,033
Men per woman over 75	0.75	0.60
	0.73	0.36
Female marriage rate over 75	NA	0.24
		0.05
States with CON law <sup>b</sup>	50	40
States with bed moratorium <sup>b</sup>	6	12
Share of Medicaid bed-days <sup>c</sup>	0.29	0.70
	0.22	0.10

Notes: NA indicates "not available." Where applicable, standard deviations appear below means.

Sources: NCSS (1974); HCIA (1996); U.S. Department of Health and Human Services (1996); Harrington et al. (1997).

<sup>a</sup> Defined as number of married women divided by total number of women. Data were obtained by the authors from the U.S. Bureau of the Census and are available upon request.

<sup>b</sup> Refers to total number of states. Data for 1971 represent 1982 values.

<sup>c</sup> Refers to number of Medicaid bed-days divided by total bed-days.

$< 0$  hold up well.  $\beta_1$  hovers near 0.9 and always remains strictly below unity. Since the point estimates always satisfy  $0 < \beta_1 < 1$  and  $\beta_2 < 0$ , one can never reject these hypotheses. More strongly, as shown in the table, in four of the six cases analyzed, one can use a one-tailed test to reject the alternative hypothesis that  $\beta_1 \geq 1$  with 99-percent confidence. Using a similar one-tailed test, one can reject the alternative hypothesis that  $\beta_2 \geq 0$  with 99-percent confidence in every case. A one-percentage-point increase in the sex ratio among people over 75 lowers the nursing home population by about 0.4 percent. The effect goes up to about 0.7 percent for the over 65 population. This seems sensible, because younger men are healthier and should be better able to care for sick mates. Moreover, the government policy variables, which we have chosen not to model explicitly, account for almost none of the observed within-county,

<sup>17</sup> Harrington et al. (1997) present data for 1978, 1982, 1986, 1990, and 1994. We assume that the series does not change from 1990 to 1991 and use the 1990 values to proxy the 1991 data. This seems reasonable, because there is very little movement in the reported data from 1990 to 1994.

<sup>18</sup> See Harrington et al. (1997).

TABLE 2—DETERMINANTS OF LONG-TERM CARE IN U.S. COUNTIES (1971–1991)

	Over 65		Over 75			
Elderly population	0.86 <sup>†</sup>	0.75 <sup>†</sup>	0.95	0.92 <sup>†</sup>	0.81 <sup>†</sup>	0.97
	29.97	24.26	17.53	32.47	26.79	18.57
Ratio of males to females	-0.80 <sup>‡</sup>	-0.78 <sup>‡</sup>	-0.89 <sup>‡</sup>	-0.40 <sup>‡</sup>	-0.32 <sup>‡</sup>	-0.59 <sup>‡</sup>
	-8.13	-7.70	-5.12	-7.34	-5.60	-4.59
Share of Medicaid bed-days <sup>a</sup>		0.34*	0.59*		0.32*	0.55*
		7.59	7.26		7.25	6.86
Certificate of Need law			0.08*			0.07*
			5.90			5.48
Nursing home bed moratorium			-0.04*			-0.06*
			-2.87			-3.70
R <sup>2</sup>	0.26	0.27	0.27	0.27	0.27	0.27
Observations	21,872	17,789	9,154	21,872	17,789	9,154

Notes: *t*-statistics are given below point estimates. The dependent variable is the log of county nursing home residents. All regressions are run within counties and years.

<sup>†</sup> Less than or equal to unity with 99-percent confidence.

<sup>‡</sup> Less than zero with 99-percent confidence.

\* Significantly different from zero with 99-percent confidence.

<sup>a</sup> Refers to statewide share.

TABLE 3—DETERMINANTS OF LONG-TERM CARE CHANGES IN U.S. COUNTIES (1971–1991)

	Over 65		Over 75	
Elderly population	0.94	0.81 <sup>†</sup>	0.97	0.94
	14.50	11.50	15.65	13.15
Male-female ratio	-0.72 <sup>‡</sup>	-0.63 <sup>‡</sup>	-0.35 <sup>‡</sup>	-0.33 <sup>‡</sup>
	-3.74	-3.10	-3.17	-2.54
Share of Medicaid bed-days <sup>a</sup>		0.21*		-0.03
		1.98		-0.82
R <sup>2</sup>	0.03	0.02	0.03	0.03
Observations	18,634	14,718	18,634	14,718

Notes: *t*-statistics are given below point estimates. The dependent variable is the log difference of countywide nursing home residents. Independent variables are also first-differenced, and year dummies are used.

<sup>†</sup> Less than or equal to unity with 99-percent confidence.

<sup>‡</sup> Less than zero with 99-percent confidence.

\* Significantly different from zero with 95-percent confidence.

<sup>a</sup> Refers to statewide share.

within-year variation.<sup>19</sup> In addition, several potentially confounding variables, such as countywide real per capita income or the countywide share of individuals from 45 to 60 (which mea-

<sup>19</sup> The initially puzzling positive coefficient on the CON law variable probably owes itself to endogenous government policy: states impose CON laws when their nursing home bed use becomes relatively high.

sures the availability of child caregivers), were found to be insignificant.

The results of the second method, in which we first-difference the data and include year dummies, are reported in Table 3.<sup>20</sup> The coefficients are quite similar in magnitude to those produced by the first method, although the fit worsens. Once again, the sex ratio coefficient  $\beta_2$  is significantly negative with 99-percent confidence in every specification. The population coefficient  $\beta_1$  is always less than unity; we cannot reject the hypothesis that  $\beta_1 \leq 1$ , although in only one of the four cases can we make the stronger claim that the hypothesis  $\beta_1 > 1$  can be rejected. First-differencing generates a different error process than the within-county estimates; the poor fit provides evidence that this error process is noisier, but the similar coefficients suggest that the noisier error process does not covary with the independent variables enough to change our estimates.<sup>21</sup>

<sup>20</sup> Unfortunately, we cannot replicate the third specification using the first-differenced data, because there is not enough consistent year-to-year variation in CON laws or bed moratoria.

<sup>21</sup> The worsening fit could owe itself to the lack of year-to-year variation in the sex ratio during the 1980's. Therefore, any model which captures the effects of the sex ratio during the 1970's will appear noisy during the 1980's, and vice versa. We should note that the fit improves if we run these regressions over the 1970's alone.

The gender effect is consistently significant statistically and economically across specifications. In all cases, relative increases in elderly males decrease per capita market output with 99-percent confidence. Moreover, the magnitudes are also quite large: for people over 65, a ten-percentage-point increase in the rate of marriage among elderly women (which can be thought of as roughly equal to the male-female ratio) decreases per capita nursing home residents by around 7 to 9 percent; for those over 75, a similar increase in the marriage rate (among elderly females) decreases per capita nursing home residents by around 4 to 5 percent.<sup>22</sup>

Tables 2 and 3 present county-level evidence consistent with our predictions for long-term care and the sex ratio. Elsewhere, we present micro-level evidence for our predictions,<sup>23</sup> where we present calculations based on a model estimated by Stern (1995). We find, in individual-level data, that spouses are willing and able to care for their frail mates; this remains true for all but the most severely disabled individuals. For most of the frail elderly population, the presence of children does not significantly reduce the importance of spousal care. Finally, we find that the aggregate implications of the individual-level estimates appear consistent in magnitude with the county-level estimates.

The estimated effect of changes in the sex ratio appears to explain most of the rising per capita output of long-term care during the 1970's. It was during this decade that the population of elderly women was growing twice as quickly as the population of elderly men. The effect of this expansion in widows is illustrated in Table 4.

We start in the first column of results with a weighted least-squares regression of county-level nursing home residents on the elderly population and set of year dummies.<sup>24</sup> Since the weights are given by the fraction of the total nursing home population resident in each

county, the coefficient on each year dummy approximately represents the growth in output, unexplained by population growth, from 1971 to the given year. Adding the sex ratio to this regression explains the majority of this unexplained growth during the 1970's. From the second column of results, we see that fully 60 percent of the growth from 1971 to 1978 is explained by changes in the sex ratio. The last column shows that an additional 20 percent of this 1970's growth is explained by growth in Medicaid subsidization.

However, observe that the sex ratio (or Medicaid, for that matter) does not explain the declines in per capita output from 1978 onwards. This should not be surprising: from 1971 to 1978, the ratio of males to females over age 75 fell from 0.64 to 0.56; after 1978, it hardly moved. Nonetheless, from 1978 to 1991, there is a 28-percent decline in the per capita output of long-term care. Roughly the same is true after controlling for the sex ratio and Medicaid. We argue that this decline represents the effect of declining disability incidence. This can be shown by examining health data at the individual level. Using an individual-level model of living arrangement choice estimated by Stern (1995), we calculated an estimated probability of nursing home entrance for different individuals, depending on age, sex, race, marital status, and disability. This allowed us to estimate the probability of nursing home entrance for different demographic groups, broken down along age, sex, race, marital status, and disability. We then used population data from the U.S. Bureau of the Census and disability data from Manton et al. (1997) to calculate the size of each demographic group.<sup>25</sup> Knowing the probability of

<sup>22</sup> In Lakdawalla and Philipson (1999), we also present cross-sectional evidence using direct estimates of the marriage rate, rather than just estimates of the sex ratio. We find that the results support our predictions.

<sup>23</sup> See Lakdawalla and Philipson (1999).

<sup>24</sup> No fixed effects are included in this weighted regression, because we want to allow the weights to change over time.

<sup>25</sup> The Census data are taken from the Census Bureau website, at [www.census.gov/population/www/estimates/usosp.html](http://www.census.gov/population/www/estimates/usosp.html). Since we do not have age-specific measures of marital status, we impute this by assuming that all available elderly men within a 10-year age-group are matched to elderly women within that group. This is not entirely satisfactory, but since there is very little change in the sex ratio during the 1980's, this simplification is unlikely to affect the results. We take from Manton et al. (1997) age-specific estimates of disability prevalence among the elderly from 1982 to 1994. This allows us to estimate the size of the disabled population. Due to limitations in the aggregate disability data over time, we are forced to assume that disability rates are constant across race and marital status.

TABLE 4—EXPLAINING AGGREGATE TRENDS IN LONG-TERM CARE

	Full-Time Trend	Sex-Adjusted Time Trend	Medicaid-Adjusted Time Trend
Population over 75	0.92*	0.90*	0.89*
	173.29	153.66	138.48
Male-female ratio over 75		-1.23*	-1.29*
		-16.20	-15.53
Share of Medicaid bed-days <sup>a</sup>			0.14**
			2.47
1971 <sup>b</sup>	0.00	0.00	0.00
1973	0.04	0.02	0.00
	1.32	0.60	0.16
1976	0.11*	0.05**	0.03
	4.00	2.01	1.10
1978	0.14*	0.06**	0.04
	4.69	2.18	1.33
1980	0.09*	-0.01	-0.05
	3.342	-0.56	-1.544
1982	0.06**	-0.05	-0.09**
	2.33	-1.71	-2.68
1986	-0.02	-0.12*	-0.16*
	-0.56	-4.30	-4.39
1991	-0.14*	-0.25	-0.32*
	-5.31	-0.95	-8.37
R <sup>2</sup>	0.93	0.94	0.94
Observations	21,872	21,872	17,789

Notes: Where applicable, robust *t*-statistics are given below point estimates. All regressions are weighted by the county's share of nationwide nursing home residents.

\* Significantly different from zero with 99-percent confidence.

\*\* Significantly different from zero with 95-percent confidence.

<sup>a</sup> Refers to statewide share.

<sup>b</sup> Excluded year dummy.

entrance for each population group, and the size of each group, we predicted overall growth in the nursing home population from 1982 (the first year of Stern's sample) onwards, and we predicted the growth that would have occurred *absent* any improvements in disability. The results are displayed in Figure 3. The heavy dashed line represents the total estimated nursing home population, while the light dashed line represents the nursing home population that would have resulted had disability remained constant. Without changes in disability, the nursing home population would have grown at roughly the rate of population; this is consistent with the evidence in Table 4. However, changes in health are estimated to have caused a substantial decline in per capita output. From 1982 to 1994, actual per capita output declined by about 13 percent. Changes in disability explain about nine percentage points, or 70 percent of this decline. The 1980's, therefore, appear to

have been dominated by disability changes, rather than changes in the sex ratio. This explains why our county-level estimates, which do not explicitly account for disability, cannot explain much of the aggregate movement in nursing home output during the 1980's. It also explains why the 1980's look so different from the 1970's, which were dominated by changes in the sex ratio.

### III. Concluding Remarks

We derived and found substantial empirical support for the surprising result that aging may be negatively related to output in long-term care markets. In particular, we found that relative increases in the share of elderly males actually drive down the output of market care and that healthy aging by the elderly may also contract it. We found that improvements in elderly health decrease the output of market care di-

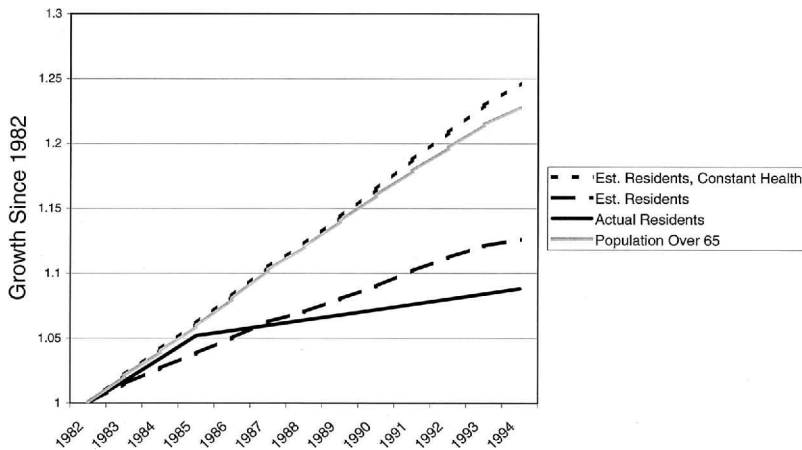


FIGURE 3. ESTIMATED GROWTH IN NURSING HOME RESIDENTS, FOR ALL AGE-GROUPS OVER THE AGE OF 65

rectly, by shrinking the base of people who need care, *and* indirectly, by raising the supply of healthy elderly people who can provide care at home. We argued that these results play an important role in explaining the changing face of aggregate market output over the past 20 years in the United States—changes that seem very surprising in light of several significant offsetting market forces. A dramatic increase in the scarcity of elderly males drove up per capita market output during the 1970's. During the 1980's, however, that scarcity did not worsen further, and health improvements resulted in substantially declining per capita market output.

Our analysis may be applied to the experiences of other developed countries undergoing demographic transitions similar to that of the United States. There is some evidence that similar patterns of rising and then declining per capita nursing home output have been observed in other developed countries. The experiences of several OECD countries, including Australia, Canada, and Sweden, appear to match the pattern illustrated in Figure 1, of rising and then declining per capita market output.<sup>26</sup> In all three countries, nursing home bed growth initially outstrips growth in the elderly population, but is eventually overtaken by population growth. These intriguing similarities suggest that further investigation into each country's experience

may reveal forces similar to those studied here. Indeed, it seems reasonable to suppose that in any economy where the elderly care for each other outside the market, long-term care markets may respond to aging in the unexpected ways we have explored. Ultimately, the aging of the population represents not just a new source of long-term care demand, but it may also represent a new source of long-term care supply, where care may be produced in the household rather than in the market.

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<sup>26</sup> These data are available from the authors upon request.

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