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

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Working Paper No. 146

**“THE RISE IN OLD AGE LONGEVITY AND THE
MARKET FOR LONG-TERM CARE”**

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

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Abstract

This paper analyzes how markets for old-age care respond to the aging of populations. We consider how the biological forces, which govern the stocks of frail and healthy persons in a population, interact with economic forces, which govern the demand for and supply of care. We argue that aging many times may lower the demand for market care by increasing the *supply* of family-provided care, which substitutes for market care. By providing healthy spouses, aging may increase the supply of family care-givers. Unexpectedly, this implies that relative growth in healthy elderly males may contract the long-term care market, while relative growth in healthy elderly females may expand that market. We examine these implications empirically using individual, county, and national-level evidence on the US market for long-term care and find substantial support for them, particularly the negative output effect of growth in elderly males. We then decompose the per capita growth in long-term care output over the last three decades into the component accounted for by improvements in health and that accounted for by relative growth among elderly males. The novel effects of unbalanced gender growth among the elderly appear important in explaining the net decline in US per-capita output over the last 30 years, a decline which seems remarkable given the simultaneous rise in demand subsidies for long-term care, declining fertility rates, rising female labor-force participation, and the deregulation of entry barriers to the nursing home industry.

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1 Introduction

The sizable reductions in fertility and mortality rates that have accompanied the demographic transition in many countries around the world have led both private and public sectors to start grappling with the care of aging populations. Since 1960, the share of the US population above 65 years of age has grown substantially, from about 9 percent to 14 percent. However, both the level and growth of this share are lower in the US than in other developed countries. For example, in many European nations, the elderly population makes up about one-fifth of the total population, and growth in this share has been larger than in the US over the past few decades. In Sweden, for instance, where the share of people above 65 makes up about one-fifth of the overall population, long-term care expenditures make up about one-third of all health care spending, compared to their one-tenth share of current health care spending in the US.¹ As growth in the older populations of developed countries has occurred, the share of public spending on the elderly accounted for by long-term care has grown as well. This 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.

Many discussions by economists seem to predict that aging will induce a rapid expansion in the market for long-term care. Implicitly, these discussions argue that because *per capita* demand for long-term care rises at an increasing rate ('exponentially') with age, growth in elderly populations swells the market for long-term care more than proportionally. These arguments imply the quantitative prediction that a one percent increase in the elderly population will result in a larger than one percent increase in the market output of long-term care. Figure 1 provides evidence relevant to the relationship between aging and long-term care growth in the US over the last few decades. The figure compares the relative growth in current nursing home residents to that in the population over the age of 75.² All series are normalized at 1971 to a value of unity.³ From Figure 1, we learn that growth in nursing home residents has *rapidly* decelerated since 1970, in spite of roughly constant rates of elderly population growth: specifically, in the mid '70s, the resident population grew at a 4.8% annual rate; by the early '80s, this annual growth rate had plummeted by almost two-thirds to 1.7%; finally, in the late '80s to early '90s, annual growth again dropped by more than three-quarters to about 0.4%. This sharp deceleration has occurred in spite of relatively stable growth rates for the elderly population: the population over 75 has grown at a roughly stable annual rate of 2.7% for the past two decades. In the 1970s, the resident population grew twice as fast as population, so that the per capita quantity of nursing home output actually grew as fast as population; over the past decade, the resident population has grown at less than half the rate of population, so that per capita quantity has fallen at half the rate of population growth.

Remarkably, this sharp decline in per capita output has taken place in spite of four significant factors which should have spurred per capita growth. First, the share of output that is publicly financed through Medicaid has grown enormously, from about 24 percent of 1971 nursing home bed-days in nursing homes to about 70 percent of bed-days by 1991.⁴ Second, from 1920 to 1940, birth rates fell by almost thirty percent;⁵ a 70 year-old in 1990 would thus have had far fewer children to provide home-produced care in *lieu* of market-based care. Third, compounding the decrease in fertility, the rapidly rising rate of female labor force participation, in concert with rising female human capital and wages, has increased the cost of family care, which is often provided by daughters. The female labor force participation rates of the 1920 cohort of children, whose parents became elderly around 1970, was 53.3 percent for women between 45 and 50; for the 1930 cohort, the rate was 61.1 percent. (Killingsworth and Heckman (1986)) Finally, towards the mid-'80s, entry and investment barriers in the nursing home industry posed by Certificate of Need legislation relaxed considerably, so that the

¹SOU (1996) *Behov Och Resurser i Vården - En Analys*, Statens Offentliga Utredningar, Stockholm.

²Most consumers in the long-term care market are above the age of 75. In 1995, about 17% of residents were 65-74, 42% between 75-85, and 41% above 85 years. (National Center for Health Statistics (1997))

³The figure uses data on nursing home residents from the 1973, 1977, 1985, and 1991 National Nursing Home Surveys, reported in Strahan (1997). Intermediate years are interpolated, and data for 1971-2 are extrapolated, assuming constant rates of growth between observed points. Population data from 1970 to 1991 comes from the National Center for Health Statistics Web Site: <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.

⁴See National Center for Social Statistics (1974) for 1971 Medicaid data. 1991 data based on HCIA (1996).

⁵See Bureau of the Census (1965), series B19-30, and B31-36.

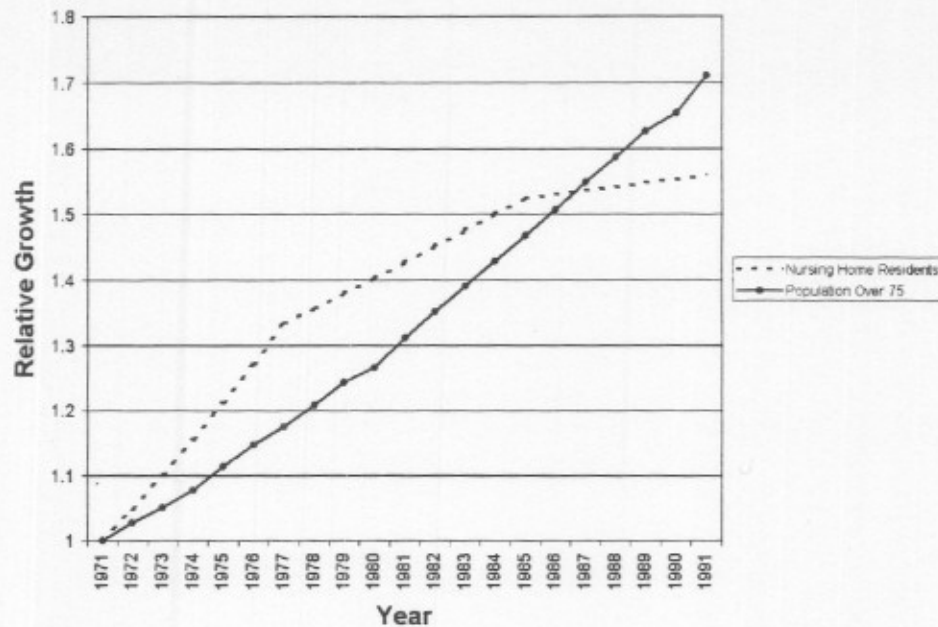


Figure 1: Relative Growth of Nationwide Nursing Home Residents *versus* Relative Growth of Elderly Population.

supply of nursing home beds should have risen and driven down the price of market-based care.⁶ It is thus quite remarkable that per capita output in this market has fallen so sharply in recent years. Per capita output contracted so sharply during the '80s that it offset the growth which occurred during the '70s; from 1971-1991, per capita output fell by almost fifteen percent overall, in spite of these four growth-inducing factors. Therefore, the central question posed by Figure 1 must be: why did per capita output change course during the '80s, and how did it contract so sharply in the presence of so many forces pushing demand higher?

Motivated by this question, our paper provides a theoretical and empirical analysis of the relationship between aging and the growth of long-term care markets. Section 2 studies the impact of aging on the equilibrium in the long-term care market. We argue that aging may actually decrease the per capita demand for long-term care; this will obtain if it raises the *supply* of family care, a primary substitute for market care, and thus reduces the demand for market care. This effect may be counteracted or reinforced by changes in frail relative to healthy life-span.

Section 3 generalizes the analysis of section 2 in order to consider the impact of aging on the supply of family care, which frail individuals may use as a substitute for market-based care. An investigation of spousal care reveals the surprising implication that market-based output *contracts* with the longevity of the scarcer sex, typically males, and expands with the *healthy* life-expectancies of the abundant sex, typically females. Care of the longer-living spouse, typically the female, is market-produced, while care for the shorter-living spouse, typically the male, is produced at home. Growth in elderly males reduces the per capita demand for market care, because it eases the relative scarcity of male care-givers in home production. Conversely, growth in elderly females exacerbates the scarcity and thus raises per capita demand. We also consider the effects of changes in the prices of market substitutes for nursing home care, such as market-based home health care. Since market-based home health care is almost always used in conjunction with family care, and since market substitutes make it cheaper for family members to supply care, their prices interact with our marriage effect. Specifically, a decrease in the price of market substitutes strengthens the marriage effect; when the price falls, the overall per capita demand for married people falls by more than it does for single people, because they are more likely to have a

⁶See Harrington *et al* (1997).

family care-giver whose cost of care provision falls enough to induce provision. Therefore, the recent expansion of subsidies to market-based home care has increased the importance of the marriage effect. Moreover, the expansion in market-based home care output suggests a corresponding rise in family care, which serves as a complement to it.

Both changes in health and changes in the sex ratio appear to be empirically significant in explaining trends in the per capita output of long-term care. Over at least the past decade, increases in the relative health of the elderly have served as a significant force slowing the rate of growth of long-term care. Specifically, in 1981, the incidence of disability among the population over 75 was 31.9%, while in 1991, this rate fell to 28.1%.⁷ Since there were about 13.5 million people over age 75 in 1991, this improvement in health resulted in there being half a million fewer disabled persons over 75 in 1991. This represents a highly significant share of the 1.5 million residents of nursing homes in 1991. (Strahan (1997)) The relationship between relative male growth and long-term care also bears on the US experience, because the past two decades have seen a dramatic amelioration of the shortage in elderly males in the US; this improvement represents a second key force contributing to the initial rise and subsequent fall in per capita demand. In the early 1970s, the male population over age 75 grew at a mere 1.7% annual rate, while females grew at a 3.4% annual rate; this force accounted for the dramatic rise in per capita output witnessed during the '70s. By the early '80s, this huge imbalance had been largely wiped out, as men were then growing at a 2.6% annual rate, while women were growing at a 2.9% rate. As the '80s progressed, the male growth rate caught up to and eventually even surpassed the female growth rate. However, during the '70s, 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.⁸ 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 per capita demand during the '70s.

Section 4 examines several implications of the model empirically and then goes on to quantify the relative contributions to long-term care growth of health improvements and unbalanced gender growth. First, using panel county-level data on nursing home residents and population, we test the prediction that increases in the share of males decrease the per capita demand for nursing home care. For some tests, the aggregate data are preferred over individual-level data, because our predictions concern the relations between aggregate stocks. We find evidence that the predicted effects are present in the data and that the effect of increases in the share of males is quite large in magnitude: a ten percentage point increase in the ratio of men per woman reduces the stock of nursing home residents by as much as nine percent. Next, we investigate whether or not the aggregate evidence on the sex ratio effect is consistent with individual-level evidence. Using microdata, we find that the presence of a spouse *more than halves* the probability of nursing home entrance. Moreover, we find that this result applies to all but the most severely mentally impaired elderly, and all elderly parents except those with children who live less than an hour away. Finally, we use national-level data to decompose the changes in per capita output into components due to health improvements and changes in marriage patterns. In particular, we find that a shift in the balance between the marriage effect and the health effect has accounted for the changing course of per capita output: during the '70s, the rapidly worsening shortage of males more than offset the improvements in health and, on balance, raised per capita output; with the advent of the '80s, the shortage of males ceased to worsen, and per capita output growth became dictated largely by health improvements, which pushed down per capita output for the remainder of the '80s and continues to push it down well into the '90s.

The paper relates to a rather scarce economic literature on the long-term care market and its relationship to aging.⁹ In particular, relatively little analytic attention has been paid by economists to the macroeconomic aspects of the market for long-term care, in general, or the impact of aging on this market, in particular. Understanding the macro-level interactions between demographic and economic forces in this market seems

⁷These rates are calculated from disability incidence estimates for people of ages 75-84, and over 85, all found in Manton *et al* (1997). These are converted into rates for people over 75 by using population data from *Vital Statistics of the United States*.

⁸All calculations were performed using male and female age-specific population data from the 1970-1991 issues of the US Department of Health and Human Services publication, *Vital Statistics of the United States*.

⁹For a review of the existing literature, see Norton (1997). For an overview of policy issues and this market, see, for example, Garber (1994), or Norton and Newhouse (1994). For the impact of nonprofit production on this market, see, for example, Gertler (1989), Gertler and Waldman (1992), or Weisbrod (1988).

important for understanding the movement of output over time and how it is affected by public intervention.

2 The Demographics and Economics of Long-Term Care

The demand for long-term care depends on both changes in longevity *per se* and the covariation between changes in longevity and changes in frailty. In this section, we develop a model which allows us to study these two components of demand growth.

We define long-term care as the care of an individual with a chronic condition. This definition leads us to consider a model in which an individual is healthy and then becomes chronically frail until death. The care of the individual in this last set of frail periods may be produced at home or in the market. We let T represent total lifetime in years, with life-expectancy given by $\mu \equiv E[T]$; S is the disabled or frail time in years, with expected value $\mu^S \equiv E[S]$; consequently, $T-S$ represents healthy years, while $\mu - \mu^S$ represents expected healthy life-span. Let (H_t, F_t) denote the stocks of healthy and frail individuals at time t . Under time-independent exit rates, these stocks change over time according to

$$\dot{H}_t = e - \frac{1}{\mu - \mu^S} H_t \quad (1)$$

$$\dot{F}_t = \frac{1}{\mu - \mu^S} H_t - \frac{1}{\mu^S} F_t, \quad (2)$$

The variable e represents the size of the entering cohort,¹⁰ and the inverse of healthy and frail life-expectancies define hazard rates by healthy people into frailty and frail people into mortality, respectively. The healthy stock is augmented by new entrants but depleted by those becoming frail. The frail stock is thus augmented by newly disabled individuals and depleted by exits due to mortality. In the steady-state, $H = e(\mu - \mu^S)$, $F = e\mu^S$, and $H + F = e\mu$.

These steady-state results generalize easily when the exit rate into disability or mortality is age-dependent. Consider the bivariate random vector (T, S) , where T represents age at death and S represents length of disabled life-span. Given a steady state population, this implies the two survival curves in Figure 2 below, where the top survival function is for the overall lifetime survival $S_T(t)$, the probability that an entrant is alive at time t , and 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 curves are constructed from estimates of US longevity and disability incidence for the US population over the age of 75.¹¹ The frail members of a cohort have completed their healthy durations, $T - S \leq t$, but not their life, $T \geq t$. Consequently, the total frail population at all ages is given by integrating across ages the unconditional probability of being alive minus the unconditional probability of being healthy. This must be equal to the patterned black area between the two curves in Figure 2 and can be expressed as:

$$F = e \int [S_T(t) - S_{T-S}(t)] dt = e\mu^S$$

Similarly, since the solid gray area below the lower $S_{T-S}(t)$ curve in Figure 2 represents the stock of healthy individuals, this stock must satisfy:

$$H \equiv e \int S_{T-S}(t) dt = e(\mu - \mu^S)$$

These stocks are identical to those generated by the model with time-independent exits.

The total demand for long-term care by frail individuals is defined by the stock of frail multiplied by their per-capita demand $d(p, \frac{H}{F})$. We assume that demand per frail person falls in the price of market care and the per

¹⁰We abstract here from the fertility effects of reductions in mortality; in other words, e is not a function of μ .

¹¹The $S_T(t)$ curve is constructed, using an assumption of a constant hazard rate into death, from expected longevity in 1991, which is 11.1 for 75 year-olds. (US Department of Health and Human Services, *Vital Statistics of the United States 1991*) The $S_{T-S}(t)$ curve is constructed using data on 1994 disability incidence for 65-74 year-olds, 75-84 year-olds, and those over 85. (Manton et al (1997)) These figures are taken to be the rates for 70, 80, and 90 year-olds, respectively. Values for intermediate years are linearly interpolated. The curve $S_{T-S}(t)$ is then just the disability incidence at time t multiplied by the probability $S_T(t)$.

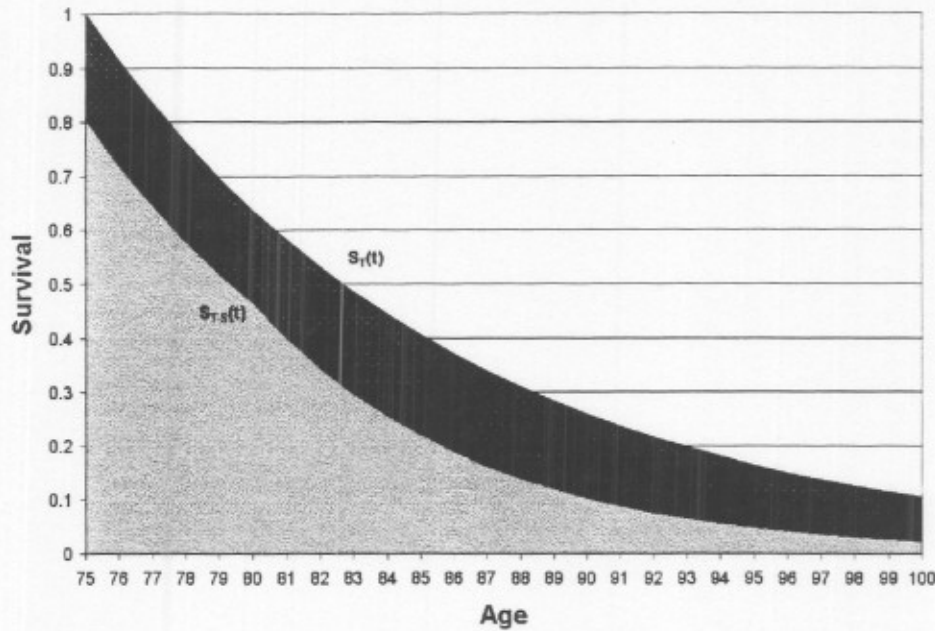


Figure 2: Overall and Healthy Lifetime Survival Functions.

capita availability of healthy care-givers: $d_p \leq 0$ and $d_H \leq 0$.¹² For now, we are also assuming that any healthy individual, regardless of family affiliation, is willing to provide home-produced care to any frail individual; this assumption is relaxed in the next section. Total demand is thus given by $D(p, H, F) = Fd(p, \frac{H}{F})$. The supply of market-based care $Z(p)$, provided by younger individuals outside our model, slopes upward due to the labor-intensity of long-term care, in which labor costs represent almost all production costs. The equilibrium quantity and the price thus satisfy:

$$Fd(p, \frac{H}{F}) = Z(p) \quad (3)$$

In this paper, we will assume that there is no technical innovation in market-based long-term care and that the function $Z(p)$ remains fixed as a result. We will thus focus on movements in demand as the key source for observed changes in output. Since the system in equations 1 and 2 may be fully characterized by the entry, e , total life-span, μ , and disabled life-span μ^S , we can express the equilibrium demand function as $D(p, e, \mu, \mu^S) = e\mu^S d(p, \frac{\mu}{\mu^S} - 1)$. Generically, define $\varepsilon_x \equiv \frac{\partial D}{\partial x} \frac{x}{D}$. This allows us to decompose growth in demand at a given price as:

$$\frac{dD}{D} \Big|_p = \frac{de}{e} \varepsilon_e + \frac{d\mu}{\mu} \varepsilon_\mu + \frac{d\mu^S}{\mu^S} \varepsilon_{\mu^S} \quad (4)$$

This decomposes demand growth into an entry component, a longevity component, and a disability component. The form of the total demand function implies the following properties for this decomposition: the elasticity of entry is unity, the elasticity with respect to longevity *per se* is negative, and the elasticity of disability is positive. Formally, $\varepsilon_e = 1$, $\varepsilon_\mu \leq 0$, and $\varepsilon_{\mu^S} \geq 0$. Since changes in longevity may be accompanied by changes

¹²Throughout the paper, we assume that the patient pays for market-based long-term care directly. Indeed, even though insurance for long-term care exists, it is little used. See Sloan and Norton (1997) for a discussion of this issue. There is, however, extensive use of public subsidies in this market. For an empirical investigation of the effects of subsidies, see Sloan, Hoerger, and Picone (1996).

in frailty, the total effect of rises in longevity must account for this effect. In particular, if $\mu^S(\mu)$ is the frail life-span associated with a given level of longevity, the total elasticity of demand with respect to longevity will be given by:

$$\frac{dD}{d\mu} \frac{\mu}{D} = \varepsilon_\mu + \varepsilon_{\mu^S} \frac{\partial \mu^S}{\partial \mu} \frac{\mu}{\mu^S}$$

The covariation between longevity and disabled life-span may reinforce or counteract the direct negative effect of longevity on nursing home demand. Declines in disability serve to reinforce this effect, while rises in disability serve to offset it.¹³

The decomposition of demand growth in equation 5 extends in a natural way to a decomposition of quantity growth. The equilibrium price and quantity may be written as a function of the three population parameters: $P(e, \mu, \mu^S)$ and $Y(e, \mu, \mu^S)$. Generically, define the elasticities of price with respect to the parameters as $\varepsilon_x^p \equiv \frac{\partial P}{\partial x} \frac{x}{P}$. It is clearly true that $\varepsilon_e^p \geq 0$, $\varepsilon_\mu^p \leq 0$, and $\varepsilon_{\mu^S}^p \geq 0$. If we define the price elasticity of demand as $\varepsilon^p \equiv \frac{\partial D}{\partial p} \frac{p}{D}$, we can now decompose total quantity growth as:

$$\frac{dY}{Y} = \frac{de}{e} (\varepsilon_e + \varepsilon_p \varepsilon_e^p) + \frac{d\mu}{\mu} (\varepsilon_\mu + \varepsilon_p \varepsilon_\mu^p) + \frac{d\mu^S}{\mu^S} (\varepsilon_{\mu^S} + \varepsilon_p \varepsilon_{\mu^S}^p)$$

As before, the growth in quantity is affected positively by entry and frailty but negatively by longevity. The price response will only partially offset the effect of the demand increase.

It is also straightforward to identify the effect of population growth and the effect of changes in per capita demand. Observe that when longevity μ and disabled life-span μ^S increase by the same percentage, the total population grows by that percentage, while the ratio $\frac{H}{F}$ remains fixed. As a result, when μ and μ^S grow by x percent, total demand grows by x percent. In other words, $\varepsilon_\mu + \varepsilon_{\mu^S} = 1$. This fact allows us to rewrite the decomposition in equation 4 as:

$$\frac{dD}{D} \Big|_p = \frac{de}{e} + \frac{d\mu}{\mu} + \left(\frac{d\mu^S}{\mu^S} - \frac{d\mu}{\mu} \right) \varepsilon_{\mu^S} \quad (5)$$

Demand growth is given by population growth, $\frac{de}{e} + \frac{d\mu}{\mu}$, along with the effect of changes in the incidence of disability $\frac{d\mu^S}{\mu^S}$, which is given by $\left(\frac{d\mu^S}{\mu^S} - \frac{d\mu}{\mu} \right) \varepsilon_{\mu^S}$. Demand growth is thus driven both by simple population growth and by the effect of changes in health.

3 Family Care & Long-term Care Markets

This section extends the analysis by investigating the novel effects of gender-biased aging on long-term care. Our model, in which patients may be cared for either by family members or by nursing homes, implies that relative increases in elderly males reduce the use of nursing homes, while relative increases in elderly females increase such use. These gender effects are strengthened when the cost to family members of providing care at home falls, as would occur when subsidies to market-based home health care aides rise, or when Social Security benefits become more generous.

Let the parameters (e_m, μ_m, μ_m^S) , denote the number of male entrants, male life-span, and male disabled life-span, respectively with the analogous set of parameters, (e_f, μ_f, μ_f^S) , for females.¹⁴ Denote by (e, μ, μ^S) the parameters for the population as a whole. Let C be the stock of healthy couples; let C_f be the stock of

¹³ Empirically, longevity μ has grown more rapidly than healthy life-span μ^S . This finding has been reported by Manton *et al* (1997) and Crimmins *et al* (1997). For example, Manton *et al* find that, holding the age distribution constant, the percentage of nondisabled persons over the age of 85 has risen, in a statistically significant way, from 34.8% in 1982 to 40.2% in 1994.

¹⁴ During the '80s, American men and women who reached age 65 had the following expected healthy and frail life-spans: $\mu_m = 15.7$, $\mu_m^S = 1.2$, $\mu_f = 21.0$, and $\mu_f^S = 3.1$. Conditional on reaching age 75, the expected life-spans become: $\mu_m = 10.1$, $\mu_m^S = 1.3$, $\mu_f = 14.0$, and $\mu_f^S = 3.1$. (Manton (1997)) Their longer life-span causes women to spend more absolute time in disability, but women also spend just under 10 percent more relative time in disability.

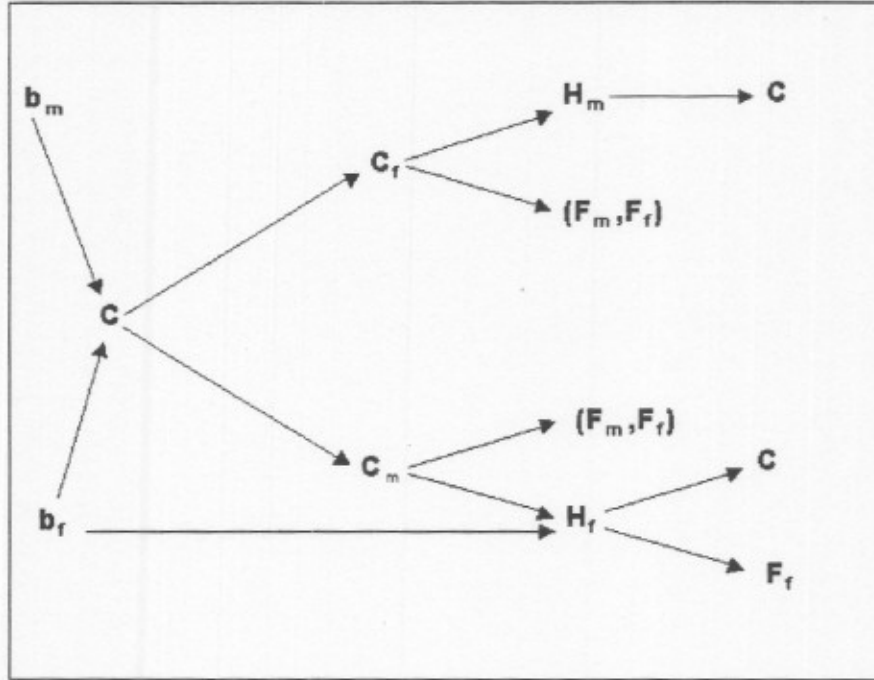


Figure 3: Transitions between Health States for Couples.

couples where the female is disabled, and let C_m be the stock of couples where the male is disabled. Denote by H_m and H_f the populations of single healthy males and females, respectively, and denote by F_m and F_f the populations of single frail males and females, respectively. We will leave the formal specification of the model to Appendix A, but we can clearly describe its dynamics in Figure 3. Healthy males and females pair off until the supply of healthy males runs out. The remaining $(e_f - e_m)$ females join the population H_f , while e_m couples join the population C . Consider a newly entering couple. Suppose that the wife falls into frailty first.¹⁵ Since her husband is still healthy, the couple joins C_f . If the wife then dies, the healthy husband is remarried to a healthy widow and rejoins the population C . If the husband falls into frailty first, however, both husband and wife join the single, frail population. Next, suppose that the husband falls into frailty first. The couple joins C_m . If the wife then falls into frailty, both mates enter the pool of unattached frail males and females, F_m and F_f . On the other hand, if the husband dies, the wife enters the pool of healthy widows, where she might be remarried into the stock of healthy couples C , or she might remain a widow until she falls into the single frail female population F_f . Observe that the abundance of healthy widows available for remarriage implies that there is no source for healthy widowers in this model. In a steady-state, therefore, there will not be any healthy unattached widowers.

The aggregate demand for nursing home care may be constructed from individual-level demands. Suppose that frail elderly can receive care from a spouse, a child, or (if no family members find it optimal to care for them) a nursing home. The demand for nursing home care by a given elderly person will thus depend on whether he is married, whether he has a child, the price of nursing home care p , and the shadow price of providing family care at home p_H . Define m as a binary indicator variable which is one if the person is married, and define x as a vector of any other characteristics affecting the care decision, such as, for example, the number of children present. We can write the frail individual's demand for nursing home care as $\delta(p, p_H | m, x)$. Clearly, δ falls in p , the price of nursing home care, and rises in p_H , the price of a substitute to nursing home care; it also falls in m ,

¹⁵To keep the model linear, we rule out the case in which both members of the couple fall into frailty at the same time, an event which occurs with probability $(\mu_m - \mu_m^S)^{-1}(\mu_f - \mu_f^S)^{-1}$ and is thus an order of magnitude less likely than the events studied here.

which indicates the availability of a spouse for the provision of informal family care. Suppose the distribution of the demand-shifters x does not differ between the married and single elderly populations.¹⁶ The per capita demand for nursing home care among frail elderly persons *who are single* may then be written as

$$d^S(p, p_H) = \int \delta(p, p_H | 0, x) d\Gamma(x).$$

The per capita demand among frail elderly persons who are married may be written similarly as

$$d^M(p, p_H) = \int \delta(p, p_H | 1, x) d\Gamma(x).$$

Since δ falls in m , aggregate demand among single people must exceed that among married people, *i.e.*, $d^S(p, p_H) > d^M(p, p_H)$. Less directly, since the average married person is more likely to have a family care-giver than the average single person, reductions in the price of home care p_H will reduce the per capita demand more among married people than among single people: $\frac{\partial d^S}{\partial p_H} < \frac{\partial d^M}{\partial p_H}$. Changes in the price of family care make no difference to individuals who have no access to family care-givers. For example, demand subsidies for home health aides (such as paid nurses, housekeepers, or others who provide health care or basic living assistance to frail individuals at home) or more generous Social Security benefits, both of which reduce the price p_H , are predicted to decrease the per capita demand among married people more than among single people.¹⁷

We assume that there is relatively more female entry into the old population, $e_m < e_f$, and that upon entrance, each healthy male is married to an entering female, so that e_m couples and $(e_f - e_m)$ single healthy females enter the aged population every period. If one partner falls into frailty before the other, the couple enters the pool of frail people with healthy partners; such frail people have the per capita demand d^M . If both partners become frail, they both enter the pool of frail people without healthy partners and with per capita demand $d^S > d^M$. Since it does not alter the theoretical results, we assume for analytical convenience that each period all healthy widows are matched to healthy widowers until the supply of healthy widowers runs out. Market output is now given by the intersection of the demand curve of frail individuals, with and without healthy spouses, with the supply curve of younger individuals:

$$(F_f + F_m)d^S(p; \Gamma, w) + (C_f + C_m)d^M(p; \Gamma, w) = Z(p)$$

Per capita demand continues to depend on the relative health of the population, but it now comes to depend also on the rate of marriage. To see this, observe that per capita demand is given by a weighted average of d^M and d^S , where the weights are given by the share of the frail population with and without healthy partners. All other things equal, increases in the rate of marriage will raise the proportion of the frail population with healthy partners and will thus weight d^M more heavily. Since $d^M < d^S$, such an increase in marriage has to lower overall per capita demand.

Following a strategy analogous to that employed in the previous section, we note that this system may be completely characterized by the parameters e and $\frac{e_m}{e_f}$, the rate of entry and the ratio of male to female entrants, μ and $\frac{\mu_m}{\mu_f}$, total longevity and the ratio of male to female life-span, and μ^S , disabled life-span. For the sake of tractability, we assume that $\frac{\mu^S}{\mu}$ is constant across gender, so that knowledge of μ , $\frac{\mu_m}{\mu_f}$, and μ^S is sufficient to identify all longevity parameters.¹⁸ Define the two additional elasticities: the gender entry elasticity is

¹⁶Even if x represents the number of children, this assumption is reasonable, provided that few elderly people have never been married.

¹⁷Implicitly, we assume that home health aides are, for the most part, complementary with family care. Individuals frail enough to require home health aides for a chronic problem are unlikely to be able to live without family assistance as well. In fact, Ettner (1984) shows that the vast majority of home health care users also receive care from family members; see the discussion in Section 4.2.

¹⁸We could also consider the independent effect of $\frac{\mu_m^S}{\mu_f^S}$, which corresponds to changes in the ratio of men to women among the frail. To keep our model manageable, however, we leave this effect to future research. It is nonetheless worth noting that Manton (1997) has found evidence that women spend a larger proportion of their elderly lives in disability, and that this effect increases with age. Under this result, aging would exert a negative effect on $\frac{\mu_m^S}{\mu_f^S}$, all other things equal.

$\varepsilon_{eG} \equiv \frac{\partial D}{\partial(\frac{e_m}{e_f})} \frac{1}{D}$, and the gender longevity elasticity is $\varepsilon_{\mu G} \equiv \frac{\partial D}{\partial(\frac{\mu_m}{\mu_f})} \frac{1}{D}$.¹⁹ We can then decompose growth in total demand, $D(p, e, \frac{e_m}{e_f}, \mu, \frac{\mu_m}{\mu_f}, \mu^S)$, as:

$$\frac{dD}{D}|_p = \frac{de}{e} + \frac{d\mu}{\mu} + \left(\frac{d\mu^S}{\mu^S} - \frac{d\mu}{\mu}\right)\varepsilon_{\mu^S} + d\left(\frac{e_m}{e_f}\right)\varepsilon_{eG} + d\left(\frac{\mu_m}{\mu_f}\right)\varepsilon_{\mu G} \quad (6)$$

There are now two components of per capita demand growth: the disability component, $(\frac{d\mu^S}{\mu^S} - \frac{d\mu}{\mu})\varepsilon_{\mu^S}$, and the gender component, $d\left(\frac{e_m}{e_f}\right)\varepsilon_{eG} + d\left(\frac{\mu_m}{\mu_f}\right)\varepsilon_{\mu G}$.

It continues to be the case that $\varepsilon_e = \varepsilon_\mu + \varepsilon_{\mu^S} = 1$, and $\varepsilon_{\mu^S} \geq 0$. While we leave the details to Appendix A, we illustrate this point with a simple example. Suppose we begin with 100 women, 90 men, and a proportion of disabled life-span $\frac{\mu^S}{\mu} = \frac{1}{5}$, which translates into a 20% frailty proportion among each sex. For simplicity, suppose there are no unattached frail males. Therefore, we begin with 10 single women, two of whom are frail, and 90 married women, 18 of whom are frail. Since we also have 18 frail married men, total market demand is just $2d^S + 36d^M$. Now suppose that e doubles, and all other parameters are held constant. Since the sex and frailty ratios are unaltered, we exactly replicate the original population, and thus we must end up doubling the population of frail singles. Specifically, we now have 200 women, 180 men, and 20 unattached women. Since the frailty ratio remains the same, 4 of the 20 unattached women will be frail, while there will now be 36 frail married men and 36 frail married women. Market demand moves to $4d^S + 72d^M$; notice that (holding constant p, p_H , and Γ) it exactly doubles. Now suppose that μ remains fixed, but that $\frac{\mu^S}{\mu}$ moves to $\frac{1}{10}$, which translates into a 10% frailty proportion. We have 1 single frail women, and 9 each of married frail men and women. Market demand, $d^S + 18d^M$, drops by half, even though population is unchanged. This demonstrates that $\varepsilon_{\mu^S} \geq 0$.

In the context of gender-specific aging, we can characterize the gender components of nursing home output growth as follows. Subject to a weak regularity condition, increases in male entry e_m or longevity μ_m lower the number of single frail females F_f . Conversely, increases in female entry e_f or longevity μ_f raise the number of single frail females. relative to female entry e_f , lower the number of single frail females F_f . This also leads to an implication for the sex ratio among the elderly: increases in relative male entry $\frac{e_m}{e_f}$ or longevity $\frac{\mu_m}{\mu_f}$ lower the number of single frail females.²⁰ As a result, if F_m and d^M are sufficiently close to zero, so that the bulk of nursing home demand comes from single frail females, $\varepsilon_{eG} \leq 0$ and $\varepsilon_{\mu G} \leq 0$.

This is a somewhat surprising and extremely important implication for projecting growth trends in long-term care. Since men are scarce as home care-givers, increases in the share of men reduce the demand for long-term care, because such increases ease the scarcity of men. Conversely, increases in the share of women exacerbate the scarcity and raise the demand for long-term care. *Note that these gender effects are over and above the effect of changes in the total number of elderly persons.* This result can be illustrated with a simple example. Suppose we have 100 women, 90 men, and a proportion of disabled life-span $\frac{1}{10}$. Initially, demand for long-term care is $d^S + 18d^M$. Suppose that either the relative number of male entrants $\frac{e_m}{e_f}$ or relative male longevity $\frac{\mu_m}{\mu_f}$ rises to equate $e_m\mu_m$ with $e_f\mu_f$. We now have 95 of each sex. We no longer have any single people, and demand is just $19d^M$. Since $d^M < d^S$, it is clear that demand has fallen, while population and the health ratio have remained the same. Increases in the share of elderly males will depress the demand for market care.

Provided that single frail females make up the bulk of nursing home patients, it is also true that growth in the population of elderly females raises demand more than proportionately. Staying with the example given above, suppose that either the number of female entrants e_f or female longevity μ_f rise by $\frac{1}{10}$, so that we have 10 more women. All of these women will be single, and one will be frail. Since only single frail people demand long-term care, the number of people demanding long-term care doubles, even though the number of women

¹⁹Note that, technically, $\varepsilon_{\mu G}$ and ε_{eG} are partial elasticities, because they refer to the percent change in demand induced by a given percentage point change in the sex ratio.

²⁰Stated formally, we have the claims that $\frac{\partial F_f}{\partial e_m} \leq 0$, $\frac{\partial F_f}{\partial e_f} \geq 0$, and $\frac{\partial F_f}{\partial \frac{e_m}{e_f}} \leq 0$. Under the regularity condition that $\mu_f^S < \mu_m - \mu_m^S$, we also have the results that $\frac{\partial F_f}{\partial \mu_m} \leq 0$, $\frac{\partial F_f}{\partial \mu_f} \geq 0$, and $\frac{\partial F_f}{\partial \frac{\mu_m}{\mu_f}} \leq 0$. The relatively weak regularity condition that $\mu_f^S < \mu_m - \mu_m^S$ holds in the US for the populations over age 65, 75, 85, 95, and 105. (Manton (1997)) All these results are proven in Appendix A.

grew by only ten percent. To understand this result, observe that simple population growth which leaves the sex ratio and the proportion of disabled life-span fixed raises population exactly proportionately: for example, replicating the elderly population will exactly double the demand for nursing home care. This implies that $\varepsilon_m + \varepsilon_f = 1$, where ε_m and ε_f denote the percent change in demand induced by a one percent change in males and females, respectively. Since growth in elderly males depresses demand, $\varepsilon_m < 0$. Therefore, $\varepsilon_f > 1$; the effect of growth in females must be disproportionate in order to balance the negative effect of growth in males.

To complete the generalization of the model, we now decompose growth in total quantity. The equilibrium price and quantity may now be written as a function of the five population parameters: $P(e, \mu, \mu^S, \frac{\varepsilon_m}{\varepsilon_f}, \frac{\mu_m}{\mu_f})$ and $Y(e, \mu, \mu^S, \frac{\varepsilon_m}{\varepsilon_f}, \frac{\mu_m}{\mu_f})$. Defining ε_{eG}^p and $\varepsilon_{\mu G}^p$ in the usual way, we know that $\varepsilon_{eG}^p < 0$ and $\varepsilon_{\mu G}^p < 0$: increases in the share of males among entrants or in the share of male longevity will lower price by reducing market demand. Since the price response only partially offsets the effect of a demand increase, we have the following two predictions:

$$\varepsilon_{eG} + \varepsilon_p \varepsilon_{eG}^p \leq 0 \quad (7)$$

$$\varepsilon_{\mu G} + \varepsilon_p \varepsilon_{\mu G}^p \leq 0 \quad (8)$$

Even after the price adjustment, increases in the ratio of males to females among entrants or the population will have negative effects on long-term care demand.

Finally, we also have the prediction that the price of home care provided by family members *interacts* with the effect of changes in the ratio of men to women; when the price of home care falls, if for example Social Security subsidies to family members rise or subsidies for home health aides rise, the gender effects, ε_{eG} and $\varepsilon_{\mu G}$, rise in absolute value. To see this, recall that increases in the price of home care raise per capita nursing home demand more for married people than for single people, or $\frac{\partial d^U}{\partial p_H} < \frac{\partial d^M}{\partial p_H}$. Increases in the ratio of men to women raise the number of married people, but lower the number of single people. As a result, the effect of an increase in this sex ratio on overall per capita demand is given by $d^M - d^U < 0$. It is thus apparent that a fall in the price of home care p_H will raise the absolute value $|d^M - d^U|$. Therefore, $\frac{\partial |\varepsilon_{eG}|}{\partial p_H} > 0$ and $\frac{\partial |\varepsilon_{\mu G}|}{\partial p_H} > 0$.

4 Empirical Analysis

In this section, we present empirical results designed to test our model and provide an explanation for the slowdown in the per capita output of long-term care. Almost all the existing empirical literature focuses on effects at the individual, rather than aggregate, level. Since ours is a model of the effects of aggregate aging implied by individual effects, we consequently use aggregate data, although we will also ensure that our results are consistent with individual data. We begin with evidence designed to test the aggregate predictions of the model. Using county-level data, we find that the aggregate share of males has a negative and extremely significant relationship with the number of nursing home residents, and that the population effect is just below unity, as predicted. *Note that this exercise could have falsified the gender-related predictions of the model, if we found no relationship between the share of males and nursing home demand.* As a result, it represents an important and relevant empirical test. Next, we use individual-level data to ensure that marriage plays an important role in nursing home entrance, above and beyond the disability status of or number of children possessed by an individual. We find that being married has a large effect; with few exceptions, it more than halves the probability of nursing home entrance, regardless of an individual's disability status or number of children. Finally, we use the theoretical model to decompose past growth in the demand for long-term care. Here we find that improvements in health and changes in the aggregate share of elderly males account for nearly all the variation in per capita demand. The rising per capita demand of the '70s owes itself to sharp increases in the share of elderly females, increases which almost entirely disappeared during the '80s. In the absence of a gender effect, improvements in health took over and drove down per capita demand. Today, the health effect continues to dominate and per capita demand continues to fall.

4.1 Assessing the Aggregate Predictions for Long-Term Care

In this section, we will test the implications we derived for the elasticity of long-term care with respect to population and gender. Unfortunately, available data do not permit an analogous test for the health elasticity, although the population and gender effects can be identified apart from the health effect under reasonable conditions.

Using the county as our unit of analysis, we will estimate panel regressions of the following form:

$$\ln(Q_{it}) = \lambda_0 + \lambda_1 \ln(\text{Pop}_{it}) + \lambda_2 \left(\frac{\text{Males}_{it}}{\text{Females}_{it}} \right) + \lambda_3 G_{it} + \phi_i + \eta_t + \varepsilon_{it}$$

The variable Pop_{it} represents the elderly (over 65 or 75) population of county i at time t .²¹ Similarly, Males_{it} and Females_{it} represent the number of males and females in county i at time t . Finally, G_{it} represents a vector of variables characterizing government regulation of the nursing home industry in county i at time t . There is also a county-specific fixed effect ϕ_i , and a year-specific aggregate effect η_t . We assume that the health effect changes over time, but not across counties within the same year. Therefore, the health effect is absorbed by η_t , which also absorbs the effect of nationwide changes in Social Security subsidies, and the effect of nationwide changes in child survival rates. It is not possible to reliably identify the effects of the latter two variables, because both change rather monotonically over time. There are two ways to estimate a model of this form; 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 aggregate effects.

It is important to note that the gender coefficient will be unaffected even if the assumption that relative health changes are uniform across counties is violated. Changes in health alone which do not have a gender component can affect only the population coefficient, but not the gender coefficient. Observe initially that only *changes* in elderly health can affect our estimates, because levels of health in each county will be absorbed by the county fixed effect ϕ_i . Moreover, if health changes are roughly uniform across counties, the impact will be absorbed by the year-specific fixed effect η_t . Therefore, suppose that in County A, μ^S falls by a fixed percentage, and that in County B μ^S rises by a fixed percentage. In this case, the gender coefficient will be unaffected, because changes in health which are uniform across genders do not affect the sex-ratio. However, this case may bias the population coefficient downward, because nursing home demand remains unchanged in both counties. The gender coefficient will be biased only if changes in unobserved total health is somehow correlated with the relative change in health across gender. In other words, there will be spurious correlation if counties which are becoming healthier tend to see healthy life-span rise by a larger *percentage* for males than for females. We know of no evidence which would support this prediction. Therefore, a positive gender coefficient is likely to reflect the positive effect of exogenous increases in the share of men. Over this period of time, such exogenous increases would have been generated by male casualties in World War I, sustained by the cohorts turning 75 in the early to middle '70s.

As a practical matter, it is difficult to separate entry elasticities from longevity elasticities, because there are no reliable longevity estimates at the county level. However, since the combined growth in entry and longevity is equivalent to growth in stocks, it is possible to identify their combined effect. In this set of regressions, λ_1 represents the elasticity of long-term care output with respect to the stock of elderly, holding fixed the sex ratio and the ratio $\frac{\mu^S}{\mu}$. Holding these two ratios fixed, the elasticity of demand with respect to the elderly stock is unity: increases in the elderly population which do not affect the share of frail or the share of men induce exactly proportionate increases in demand. Since the price response to population growth will partially offset the increase in demand, we must then have $0 \leq \lambda_1 \leq 1$. λ_2 represents the elasticity of long-term care with respect to the sex ratio. Since a given percentage change in the sex ratio is simply equal to $d\left(\frac{\varepsilon_m}{\varepsilon_f}\right) + d\left(\frac{\mu_m}{\mu_f}\right)$, the growth in gender entry plus the growth in gender longevity, λ_2 will be the weighted average of the two

²¹ 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 Pop_{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.

elasticities $\varepsilon_{eG} + \varepsilon_p \varepsilon_{eG}^p$, and $\varepsilon_{\mu G} + \varepsilon_p \varepsilon_{\mu G}^p$, where the weights are given by their respective shares in total sex ratio growth.²² Since inequalities 7 and 8 state that $\varepsilon_{eG} + \varepsilon_p \varepsilon_{eG}^p$ and $\varepsilon_{\mu G} + \varepsilon_p \varepsilon_{\mu G}^p$ are negative, λ_2 is predicted to be negative. We also have the prediction that the absolute value of λ_2 is inversely related to the price of home health care p_H : previously we saw that the absolute values of ε_{eG} and $\varepsilon_{\mu G}$ fall with p_H , so the result follows as long as a rise in p_H does not create too large an offsetting increase in the absolute value of $\varepsilon_p \varepsilon_x^p$. Unfortunately, the lack of reliable data on p_H at the county level prevents us from testing this prediction, a task we must leave to future work. We will, however, be able to adduce some suggestive evidence even without good price data.

The data we use for the county-level analysis come from the Area Resource File. (US 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 US. 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 (but not substantially) 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 notwithstanding, our qualitative results do not depend on the years used. The demographic data on elderly males and females can be broken down into males and females over the age of 65 and over the age 75. These data come from census data compiled for 1970, 1980, and 1990. To make them comparable with the long-term care data, these data were exponentially interpolated by county to construct series for 1971, 1973, 1976, 1978, 1980, 1982, 1986, and 1991. The main policy variable 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 into an estimate of total bed-days in 1971 from the Area Resource File, an estimate constructed by multiplying the number of statewide nursing home residents by 365. Intermediate years are then linearly interpolated. These data are summarized in Table 1.

INSERT TABLE 1 HERE

Note the significant decrease in the number of elderly men per woman over this 20-year period. In addition to the data reported in Table 1, 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).²³ These represent additional components of the vector G_{it} .

The results of the panel regressions are reported in Table 2.

INSERT TABLE 2 HERE

The results for both methods of estimation are reported in this table. The first method, within-county and within-year estimation, is reported in the first six columns of the table. In the first three columns of Table 2, we measure Pop_{it} and $\frac{Males_{it}}{Females_{it}}$ for the population over 65, while in the next three we measure these quantities for the population over 75. Notice that within each age group, we examine three different specifications, corresponding to different assumptions about the effect of G_{it} : in the first specification, we assume $\lambda_3 = 0$; in the second, we assume that G_{it} includes only the share of Medicaid bed-days; in the last, we assume that G_{it} includes the share of Medicaid bed-days as well as statewide CON and moratorium legislation. Regardless of how we measure the elderly population, and regardless of the assumptions about government policy, our predictions that $0 \leq \lambda_1 \leq 1$ and $\lambda_2 < 0$ hold up well. λ_1 hovers near 0.9 and always remains strictly below unity. Clearly, since the point estimates always satisfy $0 \leq \lambda_1 \leq 1$ and $\lambda_2 < 0$, one can never reject these hypotheses. More strongly, using a one-tailed test, one can reject the hypothesis that $\lambda_1 > 1$ with 99% confidence in all but

²²Formally, $\lambda_2 = d\left(\frac{\Sigma m}{\varepsilon_f}\right)(\varepsilon_{eG} + \varepsilon_p \varepsilon_{eG}^p) + d\left(\frac{\Sigma m}{\mu_f}\right)(\varepsilon_{\mu G} + \varepsilon_p \varepsilon_{\mu G}^p)$.

²³Harrington *et al* 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.

the third and sixth columns. Using a similar one-tailed test, one can reject the hypothesis that $\lambda_2 \geq 0$ with 99% confidence in every case. The data appears quite consistent with the predictions of the model. Moreover, the government policy variables, which we have chosen not to model explicitly, account for almost none of the observed within-county, within-year variation.²⁴ In addition, several potentially confounding variables, such as county-wide real per capita income or the county-wide share of individuals from 45 to 60 (which measures the availability of child care-givers), were found to be insignificant.

We should also note some evidence concerning the relationship between λ_2 and p_H : λ_2 appears to be rising over time. When the specification was estimated first for the years 1971-1980, and then separately for the years 1982-1991, the estimated λ_2 for the earlier years appeared much smaller in absolute value than for the later. We will not analyze these results, however, because they lack precision; the short time-frame of these methods accounts for the high standard errors and poor fit of these specifications. Nonetheless, this suggestive result is consistent with the increasing subsidization of home health care by Medicare and Medicaid, by which the effective price of home health care fell steadily. If this prediction is borne out by future work, the continually increasing subsidization of home health care can be seen as fueling the importance of marriage patterns in the determination of long-term care demand.

The results of the second method, in which we first-difference the data and include year dummies, are reported in the last four columns of Table 2.²⁵ The coefficients are quite similar. Once again, the sex ratio coefficient λ_2 is significantly negative with 99% confidence in every specification. The population coefficient λ_1 is always less than unity; we cannot reject the hypothesis that $\lambda_1 \leq 1$, although we cannot make the stronger claim that the hypothesis $\lambda_1 > 1$ can be rejected. This method produces an inferior fit, but the coefficient estimates are quite similar. This method generates a different error process than the first; 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 bias our estimates significantly.

The gender effect is consistently significant statistically and economically across specifications. Notice that *in all cases, relative increases in elderly males decrease per capita demand with 99% 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 seven to nine percent; for those over 75, a similar increase in the marriage rate (among elderly females) decreases per capita nursing home residents by around four to five percent. Not surprisingly, both estimation methods imply that the gender effect is stronger for the younger age group. To understand this result, consider the extreme case of a population in which expected healthy lifetime is zero, and all remaining life is expected to be spent in disability. In this case, relative increases in males have no effect, because disabled males are unable to provide home care.²⁶ The data confirm this feature of the model, since the gender elasticity is about twice as large when (presumably) healthier elderly between the ages of 65 and 74 are brought into play. Clearly, the implications of the model for gender growth and long-term care are consistent with the data. Even more importantly, however, unbalanced growth in aggregate gender stocks appears to play a quantitatively significant role in the determination of nursing home demand.

4.2 Living Arrangements and Nursing Home Care

In this section, we will present two important findings. First, spouses are willing *and* able to care for their frail mates; this remains true for all but the most severely disabled individuals. Second, the presence of children does not significantly reduce the importance of spousal care for most frail elderly. Specifically, we will calculate the probability of nursing home entrance by an individual of a specific sex, age, education, family status, and disability status; in our model, sex, age, education, having children, and severity of disability can be regarded as components of x_i . We find that the presence of a spouse decreases the probability of nursing home entrance dramatically for all except those individuals with the severest mental disabilities or children who live very close

²⁴The 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.

²⁵Unfortunately, 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.

²⁶Formally, the required regularity condition that $\mu_j^S < \mu_m - \mu_m^S$ is more likely to hold in healthier populations.

to home. In other words, spouses are able and willing to take care of frail elderly in virtually all cases. These calculations are based on estimates of a multinomial logit model of living arrangement choice found in Stern (1995).²⁷ Stern's model assumes that individuals decide whether to enter a nursing home, live without significant help from any children (this includes being cared for by a spouse alone), or receive primary care from one of their children. Due to data constraints, market-based home health care may be used by individuals reported to be in either of the latter two categories.

Table 3 provides evidence relating to the first issue, the ability of one spouse to care for another disabled spouse.

INSERT TABLE 3 HERE

The table reports the probabilities of nursing home entrance for 76 year-olds with nine years of education and no children, and various marital and disability characteristics. Across the columns, sex and marital status change, while down the rows, disability status changes. Stern's model classifies disability status by the number of Activities of Daily Living (ADLs) or Instrumental Activities of Daily Living (IADLs) individuals are unable to perform. We have grouped these ADLs and IADLs into four categories: work activities, such as cooking, doing the laundry, doing light or heavy work; movement activities, such as walking inside or outside, or traveling to the grocery store; self-care activities, such as combing or washing one's hair; and comprehension activities, such as hearing and understanding, or speaking.²⁸ The table demonstrates that, as long as a person has faculties of comprehension, spousal care is a viable and often used substitute. *For all individuals, male and female, without comprehension limitations, marriage cuts the probability of nursing home entrance by more than half.* Only for the most severely disabled people, who suffer from an inability to understand or communicate, is spousal care not an important substitute; for these people, marriage decreases the probability of nursing home entrance only slightly. In our notation, while the absolute value of δ_m falls with severity of disability, it is always negative, and it is quite large in magnitude for all but the most disabled individuals. Moreover, the most severely disabled people are not the heaviest nursing home users: Garber and MaCurdy (1990) find that moderately disabled people are the heaviest users of nursing home care, because the severely disabled tend to die fairly soon. Therefore, *the moderately disabled, who constitute the main body of nursing home residents, are highly likely to be cared for at home by a spouse, if one is present.*

These results also have important implications for market-based home health care. Since market-based home care is an option for all frail people, and since marriage has a profound impact on nursing home entrance, it cannot be true that spousal care and market-based home care are substitutable to any great extent. This result is consistent with the apparent complementarity between the two delivery systems: of the disabled elderly, 22% live in nursing homes, 21% receive family and market-based home care, but only 2% receive market-based home care alone. (Ettner (1994)) *Over ten times as many disabled elderly use market-based home care together with family care as use market-based home care alone.* Moreover, out of all patients over age 75 receiving home health care in 1994, roughly twice as many people had a primary care-giver as did not. This statistic is all the more remarkable, because roughly one-quarter of these home health care patients had no disability limitations and were thus unlikely to require a primary care-giver. These findings are consistent with our model, which treats market-based home care and family care as predominantly complementary, rather than substitutable.

INSERT TABLE 4 HERE

Table 4 provides evidence relating to the second issue, the effect of children on the importance of spousal care. This table presents probabilities of nursing home entrance and living self-sufficiently (without help from children) for a 76 year-old woman with nine years of education.²⁹ Since the probability of receiving primary care from a child is simply one minus the probabilities of nursing home entrance or self-sufficient living, it is not separately reported. Stern found that one of the most important determinants of whether a parent receives

²⁷The probabilities calculated are not found in Stern's paper, but they are based upon his instrumental variables coefficient estimates. Since some coefficients were not reported in the published paper, due to space constraints, Stern provided us with a full set of estimates.

²⁸For the definitions of the various categories, see Appendix B.

²⁹The results for men are quite similar, although the probabilities of nursing home entrance are a bit higher.

significant help from a child is, not surprisingly, the child's distance from home. Therefore, the first four rows report the effect of marriage on care arrangements when the person has 1 child living 31 to 60 minutes away, while the last four report the effect when a person has 1 child living more than an hour away, but less than a day's drive.³⁰ When the child lives more than an hour away, marriage continues to halve the probability of nursing home entrance, except for people with comprehension limitations. For people whose children live less than an hour away, marriage tends not to reduce the probability of nursing home care, but only to trade off with the probability of receiving significant help from a child. However, such people appear extremely unlikely to enter a nursing home at all: even severely disabled people with a child less than an hour away have less than a 10 percent chance of nursing home entrance. As a result, δ_m may be quite small in absolute value for people with children nearby, but such people do not make up a significant fraction of nursing home entrants. Moreover, δ_m is quite large indeed for people whose children live more than an hour away. Once again, for the main body of people at-risk for nursing home entrance, spousal care is an important and often-used substitute for market-based care.

These results are consistent with the overwhelming number of nursing home entrance studies at the individual level. For example, using similar multinomial logit frameworks, Cutler and Sheiner (1994) and Börsch-Supan *et al* (1991) find that being married significantly lowers the probability of being in a nursing home. Using a binary choice Tobit framework, Cohen *et al* (1988) find that married people are: less likely ever to enter a nursing home; likely to have fewer nursing home entrances; likely to spend fewer days per year in a nursing home. While we have focused on the effect of being married, it is also true that becoming unmarried (primarily through the death of a spouse) greatly increases the likelihood of nursing home entrance. Börsch-Supan (1990) finds that the loss of a spouse is the life event *most likely* to trigger nursing home entrance; specifically, the loss of a spouse increases *more than tenfold* the probability of transition from independent to institutional living. Moreover, Ellwood and Kane (1990) find that the loss of a spouse is the most important reason for a sudden change in income among the elderly: for men, losing a spouse decreases average annual standard of living by about ten percent; for women, the loss of a spouse decreases average annual standard of living by about twenty percent. Decreases in income, even controlling for disability, marital status, and other factors, have consistently been found to raise the likelihood of nursing home admission. (Börsch-Supan *et al* (1991))

4.3 Decomposing the Aggregate Growth in Demand

We will decompose aggregate long-term care growth from 1971 to 1991 into the population effect, and into the gender and health effects which drive changes in per capita demand.³¹ To accomplish this, we will calibrate the theoretical model presented in Section 3 to observed population growth rates, sex ratios, and disability incidence ratios,³² as reported in *Vital Statistics of the United States* and Manton *et al* (1997). Our theoretical model will be calibrated using observed estimates for total longevity, healthy longevity, and cohort size.³³ The

³⁰In a limited sample of elderly Americans, Kotlikoff and Morris (1990) found that over one-third of the elderly persons sampled either had no children or none who lived within an hour's drive.

³¹We have identified four sources of growth in per capita aggregate demand: growth in the share of females, growth in the incidence of disability, reduction in child survival, and reduction in Social Security benefits. It would not be possible to identify the latter two effects without imposing specific restrictions on the form of the individual demand function δ , as well as on the distribution of child survival. Since there is no clear way to impose such restrictions, we leave those two terms to a residual. It should be noted, however, that Social Security benefits rose from 1971-1991 and thus would have served to depress growth in per capita demand somewhat. In addition, the sharp reduction in birth rates per 1000 people from a level of 27.7 in 1920 to 19.4 in 1940 suggests that the elderly in 1970 would have had many more surviving children than those in 1990. (Bureau of the Census (1965), series B19-30 and B31-36) This may have served to raise per capita demand.

³²Using disability incidence ratios for 1982 and 1991 alone, the intermediate values are obtained by linear interpolation. Since no data are available prior to 1982, we assume that disability incidence remained constant prior to 1982. This seems a reasonably accurate approximation, since Crimmins *et al* (1997), using a much broader definition of disability, found that incidence declines during the '70s were small compared to declines during the '80s. They estimate an overall 1970 level of 42.1%, a 1980 level of 41.7%, and a 1990 level of 40.5%. (Unfortunately, the Crimmins estimates are based on a broad definition of disability which includes many health states not requiring long-term care; as a result, we cannot use these data to calibrate the model directly.) Moreover, the tremendous per capita growth in nursing home use suggests that the disability effects were more than overwhelmed by the gender effects.

³³Specifically, we take annual estimates, for populations over 65, 75, and 85, from 1971-1991, of μ , μ_f , and μ_m from *Vital Statistics of the United States*. Based on the given values of μ , and normalizing the value of e in 1970 to unity, we calibrate the

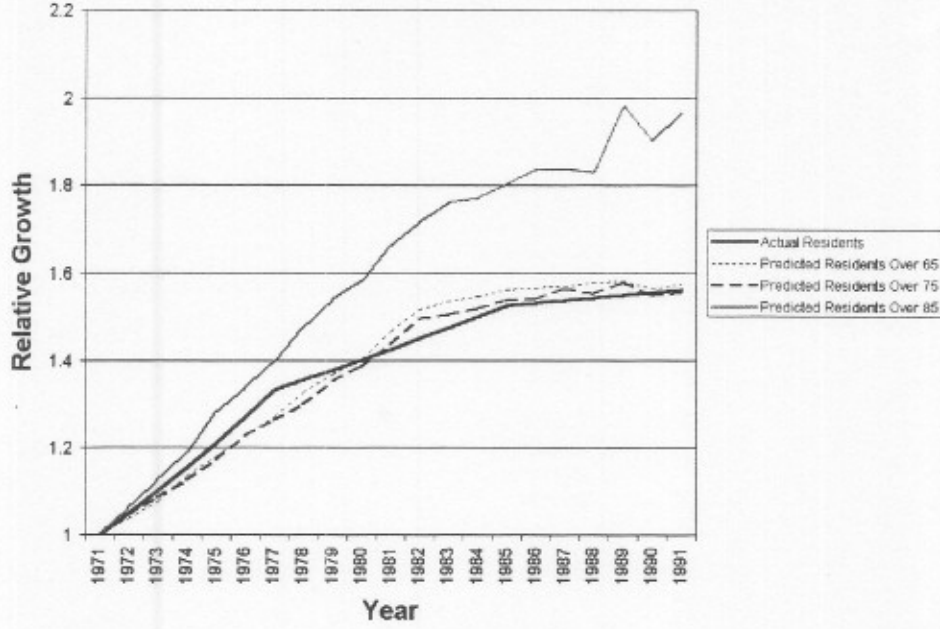


Figure 4: Actual *versus* predicted growth in US Nursing Home Residents, from 1971 to 1991.

calibrated model will then be used to predict growth in nursing home bed-days. We simplify the price response by assuming that only frail people without healthy spouses choose to buy long-term care, and that their demand is completely inelastic. The predicted growth in demand is calculated for populations over 65, 75, and 85.

Consider Figure 4, which depicts the actual growth in nursing home bed-days along with the model's predicted growth in demand for individuals over 65, over 75, and over 85. Given its simplicity, the performance of the model is quite astounding. *Predicted growth in residents over 65 and 75 tracks actual growth in residents almost perfectly*; the implications of the model are strongly consistent with the data, because the predictions of the model closely mirror actual growth in residents. Since many nursing home residents are in fact between the ages of 75 and 84, it is not too surprising that the predicted residents over the age of 85 grow much faster than actual residents. The goodness of fit should give us confidence that decomposing the growth in the predicted series will illustrate well the determinants of actual growth in nursing home care.

In our simple setting, the price-elasticity of demand is zero. Therefore, growth in predicted output \hat{Y} may be decomposed as:

$$\frac{d\hat{Y}}{\hat{Y}} = \left(\frac{de}{e} + \frac{d\mu}{\mu}\right) + \left(d\left(\frac{e_m}{e_f}\right)\varepsilon_{eG} + d\left(\frac{\mu_m}{\mu_f}\right)\varepsilon_{\mu G}\right) + \left(\frac{d\mu^S}{\mu^S} - \frac{d\mu}{\mu}\right)\varepsilon_{\mu^S}$$

The population effect is $\left(\frac{de}{e} + \frac{d\mu}{\mu}\right)$; the gender effect is $\left(d\left(\frac{e_m}{e_f}\right)\varepsilon_{eG} + d\left(\frac{\mu_m}{\mu_f}\right)\varepsilon_{\mu G}\right)$; the health effect is $\left(\frac{d\mu^S}{\mu^S} - \frac{d\mu}{\mu}\right)\varepsilon_{\mu^S}$. Given actual growth in output $\frac{dY}{Y}$, there will also be a residual effect equal to $\frac{dY}{Y} - \frac{d\hat{Y}}{\hat{Y}}$. The predicted quantities are calculated as follows. The quantity $\frac{d\hat{Y}}{\hat{Y}}$ is obtained simply by calculating the growth in beds predicted when

model to observed population growth rates g_t so that $\frac{e_t \mu_t}{e_{t-1} \mu_{t-1}} \equiv g_t$. Next, given the observed share of males m_t , we define e_{ft} and e_{mt} such that $\frac{e_{mt} \mu_{mt}}{e_{ft} \mu_{ft}} = \frac{m_t}{1-m_t}$ and $e_{mt} m_t + e_{ft} (1-m_t) = e_t$. Finally, Manton *et al* (1997) estimate disability incidence for the age intervals 65-74, 75-84, and over 85. Using population shares as weights, we construct estimates of disability incidence d_t for the populations over 65 and 75. We then calibrate μ_f^S and μ_m^S so that $\frac{\mu_f^S}{\mu_f} = \frac{\mu_m^S}{\mu_m} = d_t$.

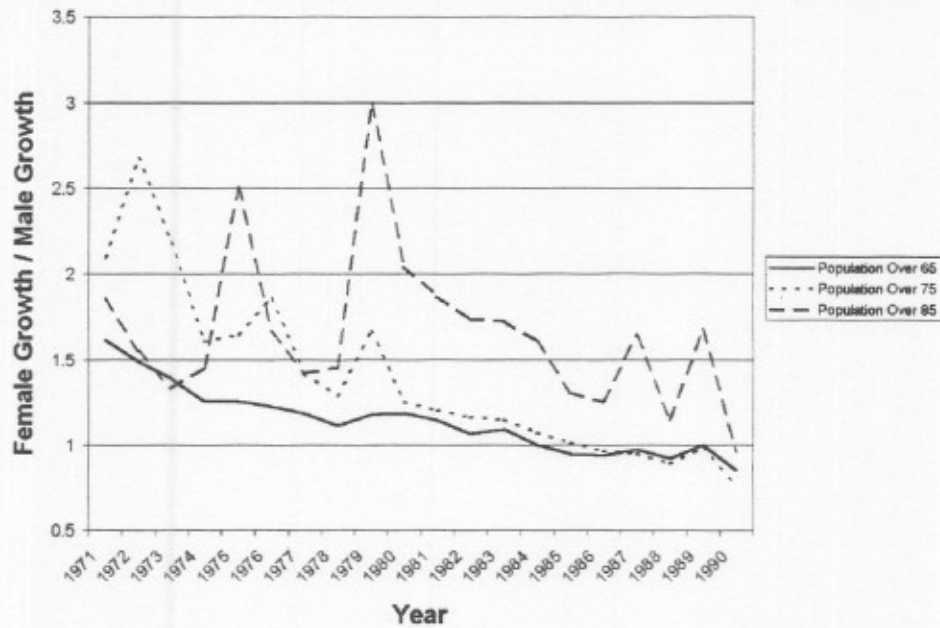


Figure 5: Growth rates of elderly females relative to growth rates of elderly males in the US.

the observed values for longevity and cohort size are used for calibration. The quantity $(\frac{de}{e} + \frac{d\mu}{\mu})$ is given simply by the total growth in the elderly population. The gender and health effects are then identified by calculating $\frac{d\hat{Y}}{\hat{Y}} - (\frac{d\mu^S}{\mu^S} - \frac{d\mu}{\mu})\varepsilon_{\mu^S}$; this is accomplished by using observed estimates for total longevity μ_f and μ_m , and cohort size e_f and e_m , but recalculating for the whole time-series the healthy and frail longevity implied by observed μ_f and μ_m and the 1980 value of $\frac{\mu^S}{\mu}$. This procedure calculates the growth predicted when relative health $\frac{\mu^S}{\mu}$ is held constant at its 1980 level, and $\frac{d\mu^S}{\mu^S} = \frac{d\mu}{\mu}$. Table 5 displays the results of this decomposition for the populations over 65, over 75, and over 85.

INSERT TABLE 5 HERE

Table 5 offers the key insight that a positive gender effect dominated growth in the '70s, but a negative health effect dominated it in the '80s. The increases in per capita output which took place during the '70s were clearly driven by a rapid increase in the number of women relative to men. During the '80s, however, this effect completely vanished. It accounted for almost no change in the over 75 age group; for the over 85 group, the gender effect actually depressed per capita demand. Figure 5, which illustrates the annual growth rate of the ratio of women to men, helps explain this change over time. During the '70s women of all elderly age groups were growing much more rapidly than men, but over time the growth rates converged toward equality. In fact, towards the late '80s, elderly men actually began to grow faster than elderly women, so that the gender effect was negative for all age groups. Table 5 also confirms our finding from the previous section that the gender effect is weaker for older, less healthy age groups. Even though the ratio of women to men is rising much faster for the over age 85 group than for the over age 65 group, the total gender effect is just slightly more than half as large. By itself, however, the near zero gender effect during the '80s does not explain the decline in per capita output, which owes itself to substantial improvements in health. Table 5 demonstrates that during the '80s, the negative health effect entirely dominated the near zero gender effects and drove down per capita output considerably. Towards the latter part of the '80s, the gender effect itself becomes slightly negative. These trends over time are demonstrated in Figure 6. The figure depicts the health and gender components of annual predicted per capita

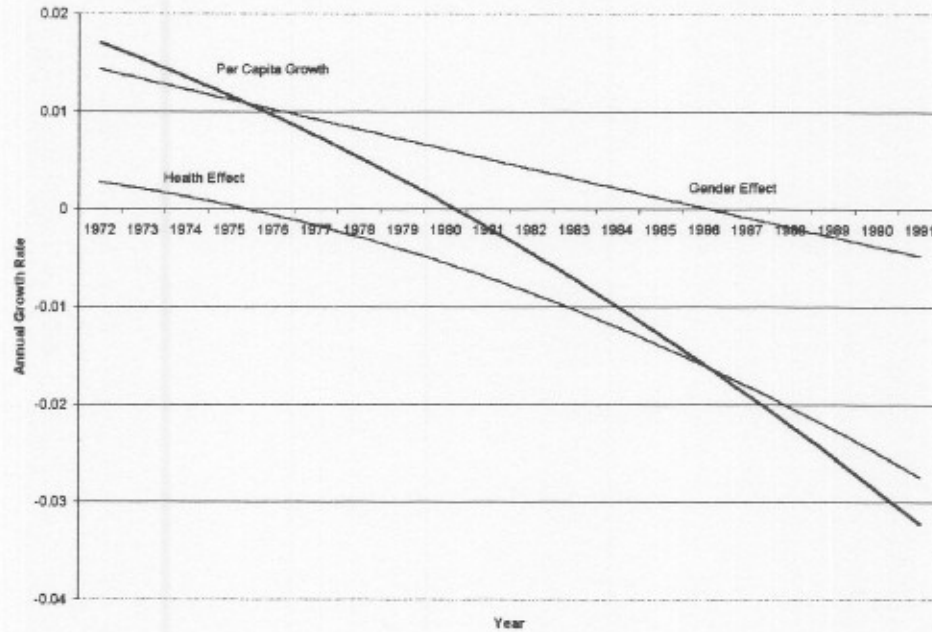


Figure 6: Decomposing Predicted Per Capita Output Growth for people over age 75.

output growth.³⁴ The gender effect does not completely vanish until the late '80s, but the negative health effect grows in magnitude and comes to dominate the movement of per capita output growth by the early '80s.

Over the past twenty years, there have been two competing effects at work in the determination of growth in per capita output: the increasing scarcity of elderly men, which has driven up per capita output, and the increasing health of elderly populations, which has driven down per capita output. During the '70s, the increasing scarcity of males overwhelmed advances in elderly health and raised the per capita output of long-term care. People were forced into nursing home care due to a lack of spousal care. The importance of the gender effect fell quickly over time, because growth in males began to catch up to growth in females. By the mid-'80s, improvements in health more than offset the barely increasing shortage of males; by the late '80s, the shortage of males actually began to ease. During this period, both gender and health forces pushed down per capita output. Given the accelerating declines in the incidence of disability,³⁵ as well as the recent balancing of gender growth, per capita output is likely to continue its decline.

5 Concluding Remarks

This paper analyzed theoretically and empirically the way in which the market for long-term care responds to aging. We identified three key forces behind movements in long-term care: elderly population growth, health improvements, and the changing sex composition of the elderly population. We derived and empirically validated the surprising result that relative increases in the share of elderly males actually drive down the demand for long-term care. This new result plays an important role in explaining the changing face of aggregate demand for long-term care over the past twenty years. Rising per capita output of long-term care during the '70s occurred as a result of a dramatic worsening in the shortage of elderly males. Per capita output fell during the '80s, because that shortage failed to worsen further, and health improvements came to dominate the movement of

³⁴All series are smoothed using a quadratic polynomial.

³⁵Manton *et al* (1997) find that the rate of disability incidence fell more quickly during the early '90s than the mid-'80s.

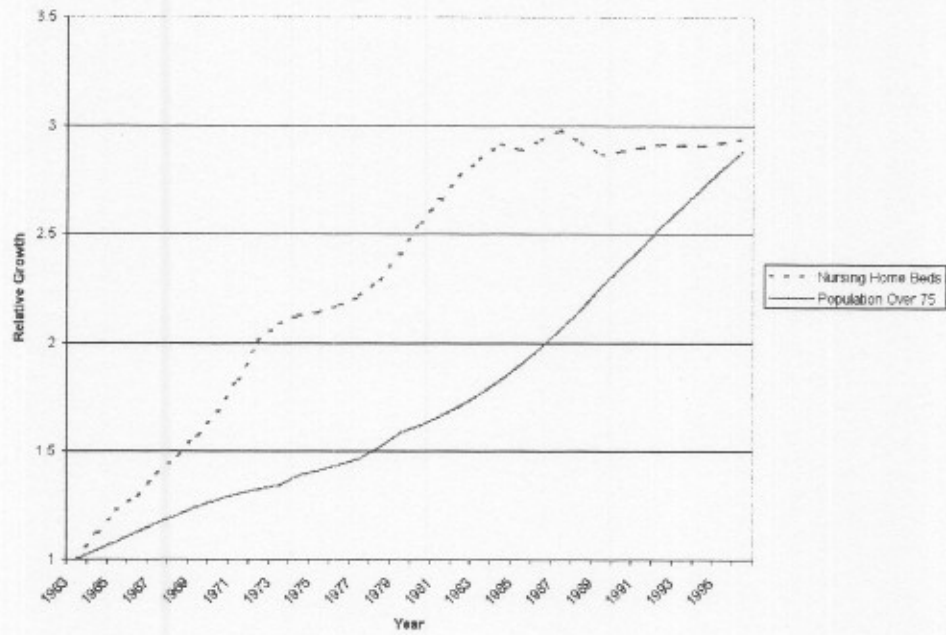


Figure 7: Relative Growth of Nursing Home Beds *versus* Elderly Population in Australia.

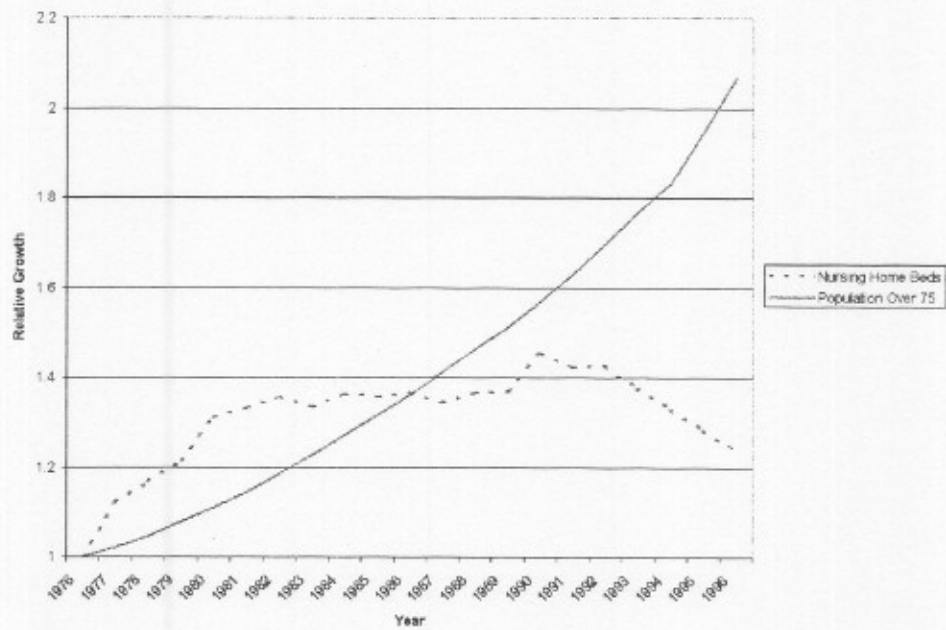


Figure 8: Relative Growth of Nursing Home Beds *versus* Elderly Population in Canada.

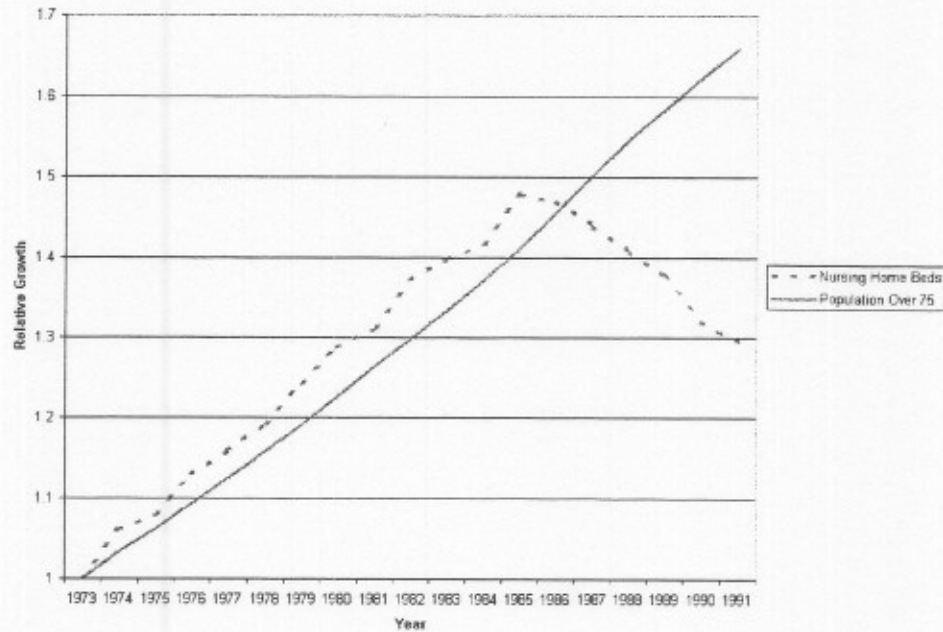


Figure 9: Relative Growth of Nursing Home Beds versus Elderly Population in Sweden

per capita demand.

Our analysis suggests two reasons why per capita demand will continue to decline over the near term. First, improvements in health among the elderly are accelerating. Second, elderly male populations are currently growing faster than elderly female populations, and if the trends over the past decade continue, the growth differential will further increase in favor of the male populations. As a result, there will be relatively more healthy elderly and relatively more elderly men to function as spousal care-givers. The declines in the per capita output of market-based long-term care evidenced in Figure 1 are thus likely to continue.

The analysis suggests several avenues of future research. First, we have stressed the importance of relative frail life-span in determining market responses to aging. While we have taken this feature as exogenously determined, explaining its movement within and across cohorts presents a more ambitious research agenda. To model the determination of frailty, one could adapt the health human capital model,³⁶ in which the health stock is augmented by health investments but depreciates at an increasing rate. In the context of the model, mortality occurs when the stock of health capital falls below a certain level. Analogously, one may interpret the onset of frailty as the time at which this stock falls below an *intermediary* level. The individual is healthy over the period during which health human capital stays above the intermediary level, and she is frail over the period during which health human capital remains between the intermediary and terminal level. If mortality gains occur for exogenous reasons over time, the longer life-spans induced will likely lead to larger investments in health. Hence, increases in longevity may increase relative healthy life-span. Therefore, an important question for future research concerns the impact of economic determinants, such as public old-age income or health support, on the relationship between frailty and mortality.³⁷ Treating frailty and mortality as economically determined would also affect the evaluation of the costs and benefits of medical research. There may be considerable costs of long-term care provision associated with medical research favoring predominantly female diseases, and considerable

³⁶See, for instance, Grossman (1972).

³⁷Of interest here would be the relationship between markets for *short-* and long-term care. Short-term health care may reduce both entry into and exit out of disability, forces which have offsetting effects on the stock of disabled. Incorporating this into a model of long-term care would raise important general equilibrium issues.

reductions in cost associated with medical research favoring predominantly male diseases. As total expenditures on long-term care grow over time, this factor may be increasingly difficult to ignore.

Second, we argued that labor-market subsidies to family care-givers, such as Social Security, lower the per capita demand for long-term care, but we took such subsidies as exogenous. In fact, they may well depend not only on the mean aging of populations but also on the *variance* of aging. Eligibility for the receipt of mandatory annuities, Social Security in the US, is age-dependent.³⁸ Interestingly, this implies that an increase in the variance of parental ages may decrease market care by making more children eligible for home-production subsidies. The variance of the parental age distribution matters, because children with relatively older parents receive more subsidies (relative to income) if they leave the labor market and provide home care for their parents.

Finally, we hope that our analysis will be applied to the experiences of other developed countries undergoing similar demographic transitions. There is some evidence that similar patterns of rising and then declining per capita nursing home output are observed in other developed countries. Figures 7, 8, and 9 depict the growth in nursing home beds and in the population over age 75 for Australia, Canada, and Sweden, respectively.³⁹ In all three countries nursing home bed growth initially outstrips growth in the elderly population, but is eventually overtaken by population growth. In the cases of Canada and Sweden, nursing home beds actually begin to decline during the late '80s to early '90s, in spite of continuing increases in the elderly population. Over the same period, Australian beds were roughly level in the face of a growing elderly population. These intriguing similarities suggest that further investigation into the specific determinants of each country's experience may reveal forces similar to those we have argued to be present in the US data.

³⁸In addition, Philipson and Becker (1997) discuss the direct impact that annuities have on longevity itself.

³⁹The data source on population over 75 and nursing home beds is OECD (1998). 1963 baseline values for Australia are: 283,200 people over 75, and 25,535 nursing home beds. 1976 baseline values for Canada are: 693,600 people over 75, and 169,700 nursing home beds. 1973 baseline values for Sweden are: 397,221 people over 75, and 34,839 nursing home beds.

A Gender Imbalance and Long-Term Care

In this appendix, we will prove claims made for the model in Section 3. To economize on notation, we will define healthy life-span as $\alpha \equiv \mu - \mu^S$, where α_m and α_f will refer to healthy life-span for males and females, respectively. In addition, we will define $\beta \equiv \mu^S$, where β_m and β_f will function similarly. The proofs will use the steady-state values from the model of marriage and long-term care described in the text, and formalized as follows:⁴⁰

$$\dot{C} = e_m - (\alpha_f^{-1} + \alpha_m^{-1})C + \beta_f^{-1}C_f$$

$$\dot{C}_f = \alpha_f^{-1}C - \beta_f^{-1}C_f - \alpha_m^{-1}C_f$$

$$\dot{C}_m = \alpha_m^{-1}C - \beta_m^{-1}C_m - \alpha_f^{-1}C_m$$

$$\dot{H}_m = -\alpha_m^{-1}H_m$$

$$\dot{H}_f = (e_f - e_m) + \beta_m^{-1}C_m - \beta_f^{-1}C_f - \alpha_f^{-1}H_f$$

$$\dot{F}_f = \alpha_f^{-1}H_f + (\alpha_m^{-1}C_f + \alpha_f^{-1}C_m) - \beta_f^{-1}F_f$$

$$\dot{F}_m = \alpha_m^{-1}H_m + (\alpha_m^{-1}C_f + \alpha_f^{-1}C_m) - \beta_m^{-1}F_m$$

All the proofs will use the following steady-state values:

$$F_m = e_m\beta_m \left(1 - \frac{\alpha_m^2}{(\alpha_m + \alpha_f)(\alpha_m + \beta_f)}\right) \left(\frac{\beta_f\alpha_m}{(\alpha_m + \beta_f)(\alpha_m + \alpha_f)} + \frac{\beta_m\alpha_f}{(\alpha_f + \beta_m)(\alpha_f + \alpha_m)}\right) \quad (9)$$

$$F_f = \beta_f(e_f - e_m) + e_m\beta_f \left(1 - \frac{\alpha_m^2}{(\alpha_m + \alpha_f)(\alpha_m + \beta_f)}\right) \left(\frac{\alpha_f}{\alpha_m + \alpha_f} + \frac{\alpha_m(\beta_f - \alpha_m)}{(\alpha_m + \beta_f)(\alpha_m + \alpha_f)}\right) \quad (10)$$

$$C_m = \frac{e_m\beta_m\alpha_f^2}{(\alpha_f + \beta_m)(\alpha_m + \alpha_f)} \left(1 - \frac{\alpha_m^2}{(\alpha_m + \alpha_f)(\alpha_m + \beta_f)}\right) \quad (11)$$

$$C_f = \frac{e_m\beta_f\alpha_m^2}{(\alpha_m + \beta_f)(\alpha_m + \alpha_f)} \left(1 - \frac{\alpha_m^2}{(\alpha_m + \beta_f)(\alpha_m + \alpha_f)}\right) \quad (12)$$

Proof of $\varepsilon_e = \varepsilon_\mu + \varepsilon_{\mu^S} = 1$

⁴⁰We abstract from the case in which both members of a couple fall into frailty at exactly the same time. Including this case would require the following changes:

$$\dot{C} = e_m - (\alpha_f^{-1} + \alpha_m^{-1} + \alpha_m^{-1}\alpha_f^{-1})C + \beta_f^{-1}C_f$$

$$\dot{H}_f = \alpha_f^{-1}H_f + \alpha_m^{-1}\alpha_f^{-1}C + (\alpha_m^{-1}C_f + \alpha_f^{-1}C_m) - \beta_f^{-1}F_f$$

$$\dot{H}_m = \alpha_m^{-1}H_m + \alpha_m^{-1}\alpha_f^{-1}C + (\alpha_m^{-1}C_f + \alpha_f^{-1}C_m) - \beta_m^{-1}F_m$$

Suppose that e_f and e_m increase by x percent. Equations 9 and 10 imply that F_f and F_m then both increase by x percent. Similarly, by inspection of equations 10, 9, 12 and 11, we can see that if α_m , α_f , β_m , and β_f all increase by x percent, F_f , F_m , C_f , and C_m all increase by x percent. Therefore, total demand increases by x percent.

QED

Proof of $\varepsilon_{\mu^S} \geq 0$

Suppose that β_m and β_f rise by x percent, and that $\frac{\alpha}{\mu}$ is constant across gender. Inspection of equations 9 and 10 reveal that F_f and F_m must rise by more than x percent. Equations 11 and 12 reveal that C_m and C_f rise by less than x percent. As a result, the rate of marriage to healthy spouses among the frail must fall. Since the total population of frail people rises, and per capita demand among the frail rises, total demand for nursing home care must rise, and $\varepsilon_{\mu^S} \geq 0$.

QED

Proof of $\frac{\partial F_f}{\partial e_m} < 0$, $\frac{\partial F_f}{\partial \mu_m} < 0$, $\frac{\partial F_f}{\partial e_f} > 0$, $\frac{\partial F_f}{\partial \mu_f} > 0$, $\frac{\partial F_f}{\partial \frac{e_m}{e_f}} < 0$, and $\frac{\partial F_f}{\partial \frac{\mu_m}{\mu_f}} < 0$.

First suppose that e_m rises. Observe that

$$\frac{\alpha_m(\beta_f - \alpha_m)}{(\alpha_m + \beta_f)(\alpha_m + \alpha_f)} < \frac{\alpha_m(\beta_f + \alpha_m)}{(\beta_f + \alpha_m)(\alpha_m + \alpha_f)} = \frac{\alpha_m}{\alpha_m + \alpha_f}$$

Therefore, the term in the second set of parentheses in 10 is strictly less than unity, and the product of the two terms in parentheses is strictly less than unity. It must then be true that $\frac{\partial F_f}{\partial e_m} = -\beta_f(1 - \xi)$, for some $\xi < 1$, and that $\frac{\partial F_f}{\partial e_m} < 1$.

Second, suppose that μ_m rises. The term in the first set of parentheses in equation 10 must fall. The condition $\beta_f < 2\alpha_m$ (implied by $\beta_f < \alpha_m$) ensures that the derivative of the numerator in the second set of parentheses is negative. This is sufficient to imply that the term in the second set of parentheses falls as well, and that $\frac{\partial F_f}{\partial \mu_m} < 0$.

It is clear from equation 10 that F_f rises in e_f . If $\frac{\alpha_m(\beta_f - \alpha_m)}{(\alpha_m + \beta_f)(\alpha_m + \alpha_f)}$ rises in μ_f , moreover, then F_f also rises in μ_f . Given the condition $\beta_f < \alpha_m$, this term rises in μ_f when the numerator is held constant. Allowing the numerator to vary only reinforces this tendency. Therefore, F_f rises in μ_f .

Now suppose F_m is at or near zero. If $\frac{e_m}{e_f}$ rises, and e stays constant, e_m rises by ϕ , while e_f falls by ϕ . The net change in F_f is then $-\beta_f\phi - \beta_f(1 - \xi)\phi < 0$. Therefore, $\varepsilon_{eG} < 0$. If $\frac{\mu_m}{\mu_f}$ rises, while μ stays constant, μ_m must rise while μ_f falls. Since F_f falls in μ_m and rises in μ_f , $\varepsilon_{\mu G} < 0$.

QED

B Disability Limitations

We group ADLs and IADLs into four categories: work limitations, movement limitations, self-care limitations, and comprehension limitations. Work limitations represent inability to perform household chores, such as cooking or laundry. Movement limitations are self-explanatory. Self-care limitations involve inability to perform tasks associated with the immediate maintenance of one's health and cleanliness, such as bathing, or washing one's hair. Finally, comprehension limitations represent inability to understand or communicate. A list of the various categories follows, in which ADLs are marked with an asterisk.

Work limitations consist of: problems cooking; problems with fatigue; problems with heavy work; problems doing laundry; problems lifting ten pounds; problems doing light work; problems reaching above one's head; problems using fingers to grasp.

Self-Care limitations consist of: problems bathing*; problems combing one's hair; problems dressing*; problems eating*; problems putting on socks; problems using the toilet; problems washing one's hair.

Movement limitations consist of: problems climbing stairs; problems getting out of bed*; problems shopping; problems with transportation; problems walking inside*; problems walking outside.

Comprehension limitations consist of: problems handling money; problems hearing or understanding; problems reading the newspaper; problems speaking; problems taking medicine; problems using the telephone.

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Table 1: Mean Attributes of Long-Term Care in US Counties (1971-91).

	1971		1991	
	Mean	Std Dev	Mean	Std Dev
Nursing Home Residents	349	1194	582	1605
Population Over 65	6699	22757	10322	31023
Men per Woman Over 65	0.81	0.23	0.71	0.09
Population Over 75	2561	8566	4365	13033
Men per Woman Over 75	0.75	0.73	0.60	0.36
Share of Medicaid Bed-Days	0.29	0.22	0.70	0.10

Source: US Department of Health and Human Services (1996).

Table 2: Determinants of Long-Term Care in U.S. Counties (1971-91).^a

	Log Nursing Home Residents ^b						Log Difference Nursing Home Residents ^c			
	Over 65			Over 75			Over 65		Over 75	
Elderly Population	0.90 †	0.78 †	0.99	0.93 †	0.82 †	0.98	0.96	0.88	1.00	0.94
	30.03	24.28	17.39	32.74	27.01	18.70	13.73	12.06	14.41	13.15
Ratio of Males to Females	-0.88 ‡	-0.86 ‡	-0.95 ‡	-0.48 ‡	-0.39 ‡	-0.68 ‡	-0.73 ‡	-0.68 ‡	-0.35 ‡	-0.33 ‡
	-8.57	-8.04	-5.23	-8.06	-6.23	-4.87	-3.59	-3.30	-2.73	-2.54
Share of Medicaid Bed-Days ^d		0.34 ††	0.59 ††		0.32 ††	0.55 ††		0.21		-0.03
		7.61	7.26		7.20	6.82		1.62		-0.82
CON Law			0.08 ††			0.07 ††				
			5.86			5.45				
Nursing Home Bed Moratorium			-0.04 ††			-0.06 ††				
			-2.81			-3.74				
R ²	0.26	0.27	0.27	0.27	0.27	0.27	0.03	0.02	0.03	0.03
Observations	21873	17790	9155	21872	17789	9154	18634	16761	18634	16939

†Less than or equal to unity with 99% confidence.

‡Less than zero with 99% confidence.

††Significantly different from zero with 99% confidence.

^aT-Statistics given below point estimates.

^bRegressions run within counties and years.

^cAll independent variables are also first-differenced, and year dummies are used.

^dRefers to statewide share.

Table 3: Disability and the Effect of Marriage on Nursing Home Entrance.^a

Disability Limitations (DLs)	Female Elder		Male Elder	
	Single	Married	Single	Married
None	0.004	0.001	0.006	0.002
Work, Movement DLs.	0.246	0.098	0.333	0.142
Work, Movement, Self-Care DLs.	0.352	0.153	0.455	0.217
Work, Movement, Self-Care, Comprehension DLs.	0.969	0.913	0.980	0.941

^aValues reflect estimated probability of nursing home entrance for a 76 year-old with no children, 9 years of education, and the additional indicated characteristics.

Table 4: Children and the Effect of Marriage on the Probability of Different Living Arrangements.^a

Distance of Child	Characteristics of Elderly Person ^b	Nursing Home		Self-Sufficient	
	Disability Status	Single	Married	Single	Married
31-60 minutes	No ADL's or IADL's	0.004	0.001	0.934	0.982
	Work, Movement DLs.	0.048	0.043	0.146	0.394
	Work, Movement, Self-Care DLs.	0.050	0.051	0.093	0.281
	Work, Movement, Self-Care, Comprehension DLs.	0.056	0.070	0.002	0.007
Over an hour	No ADL's or IADL's	0.004	0.001	0.992	0.998
	Work, Movement DLs.	0.195	0.090	0.597	0.834
	Work, Movement, Self-Care DLs.	0.256	0.136	0.470	0.751
	Work, Movement, Self-Care, Comprehension DLs.	0.475	0.518	0.015	0.049

^aThe three possible living arrangements are: nursing home residence, self-sufficient living, and living with assistance from children. The probability of the third alternative may be inferred from the others.

^bOther characteristics are held constant at their median values. The elderly person is taken to be a 76 year-old female with 9 years of education and one child. The child is taken to be the oldest female child, 47 years old, married with children, working, and with a working spouse.

Table 5: Decomposition of Growth in Long-Term Care Demand

	Over 65		Over 75		Over 85	
	1971-1981	1981-1991	1971-1981	1981-1991	1971-1981	1981-1991
Population Effect	24.32%	21.74%	27.12%	30.13%	45.73%	34.41%
Gender Effect	8.23%	1.79%	9.23%	0.33%	4.90%	-2.63%
Health Effect	5.71%	-12.16%	0.00%	-18.92%	0.00%	-10.05%
Residual Effect	-2.76%	-0.67%	-0.85%	-0.84%	-15.13%	-11.03%
Total Growth	35.50%	10.69%	35.50%	10.69%	35.50%	10.69%