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Darius Lakdawalla
Tomas Philipson

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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 and supply for labor-intensive care. Many economists have argued that aging will raise the market demand for long-term care, and hence price and quantity through classic market effects. We argue that the direct effect of aging is to lower the demand for market care by increasing the *supply* of home production. By influencing the length of frail lifetimes, aging may also have a further indirect effect, which may reinforce or counteract the direct negative demand effect. By providing healthy spouses, the marriage market provides care-givers for home production of long-term care; therefore, growth in old-age longevity may lower the demand for market production. Growth of elderly males serves to *contract* the long-term care market because it eases the scarcity of men in the old-age marriage market; growth of females serves to expand market care because it worsens the scarcity of men. These predictions lend themselves to an interpretation of the rapid deceleration in output growth that has taken place in the US over the last two decades, despite a constant rate of longevity growth and enormous *growth* in demand subsidies: since growth in elderly males has risen dramatically relative to growth in elderly females, the rate of long-term care growth has slowed significantly. We test our predictions empirically using state- and county-level evidence on the US market for long-term care in nursing homes over the last three decades. The evidence provides support for, among other things, the predictions we offer concerning the response in output growth to aging and the contraction of output due to the aging of males.

Darius Lakdawalla
Department of Economics
University of Chicago
1126 East 59th Street
Chicago, IL 60637

Tomas Philipson
Department of Economics
University of Chicago
1126 East 59th Street
Chicago, IL 60637
and NBER
t-philipson@uchicago.edu

1 Introduction

The sizable reductions in fertility and mortality rates that have accompanied the demographic transition in many countries across 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 a fifth of the total population, and growth rates in this share have 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 markets for the long-term care of older individuals, study of both the functioning of private markets for long-term care and the effects of public intervention into these markets.

Many discussions by economists seem to predict that aging will induce an rapid expansion in the market for long-term care. Implicitly, these discussions argue that because average 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 market long-term care output. This belief in the exponential growth effect draws further support from a price-elasticity of demand which seems to fall with age, as older individuals are more publicly subsidized than younger ones.

Figure 1 provides evidence relevant to the relationship between aging and long-term care growth. The figure compares the relative growth in nursing home bed-days to that in the population over the age of 75.[2] All series are normalized at 1973 to a value of unity.[3] From Figure 1, we learn that growth in nursing home bed-days has *rapidly* decelerated since 1970: in the mid '70s, bed-days grew at a 4.61% annual rate; in the mid '80s, this annual growth rate plummeted by more than half to 2.17%; finally, in the late '80s to early '90s, growth again dropped to about 1.43%. 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, while growth in the population over 65 has suffered a small drop, from 2.4% annually in the '70s to 2.05% in the '80s. In the 1970s, bed-days grew disproportionately, but over the past decade or so, they have grown much less than proportionately with the elderly population. Remarkably, this deceleration has taken place in spite of enormous growth in the Medicaid subsidization of long-term care. The share of output that is publicly financed through Medicaid has grown from about 24 percent of 1971 nursing home

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

[2] Most 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. See National Center for Health Statistics (1997).

[3] The figure uses nursing home data from 1973, 1977, 1985, and 1993, where the first three years are from the National Nursing Home Survey (1973, 1977, 1985), and the last is from HCIA (1996). Population data from 1970, 1980, and 1990, comes from the U.S. Census. Intermediate values are extrapolated exponentially, assuming a constant annual growth rate between observed points.

The 1973 baseline values are: 368,905,984 Bed-Days; 8,291,639 people over the age of 75; 3,158,456 males over the age of 75; 5,128,296 females over the age of 75. Note that males and females may not add up to total population due to extrapolation error.

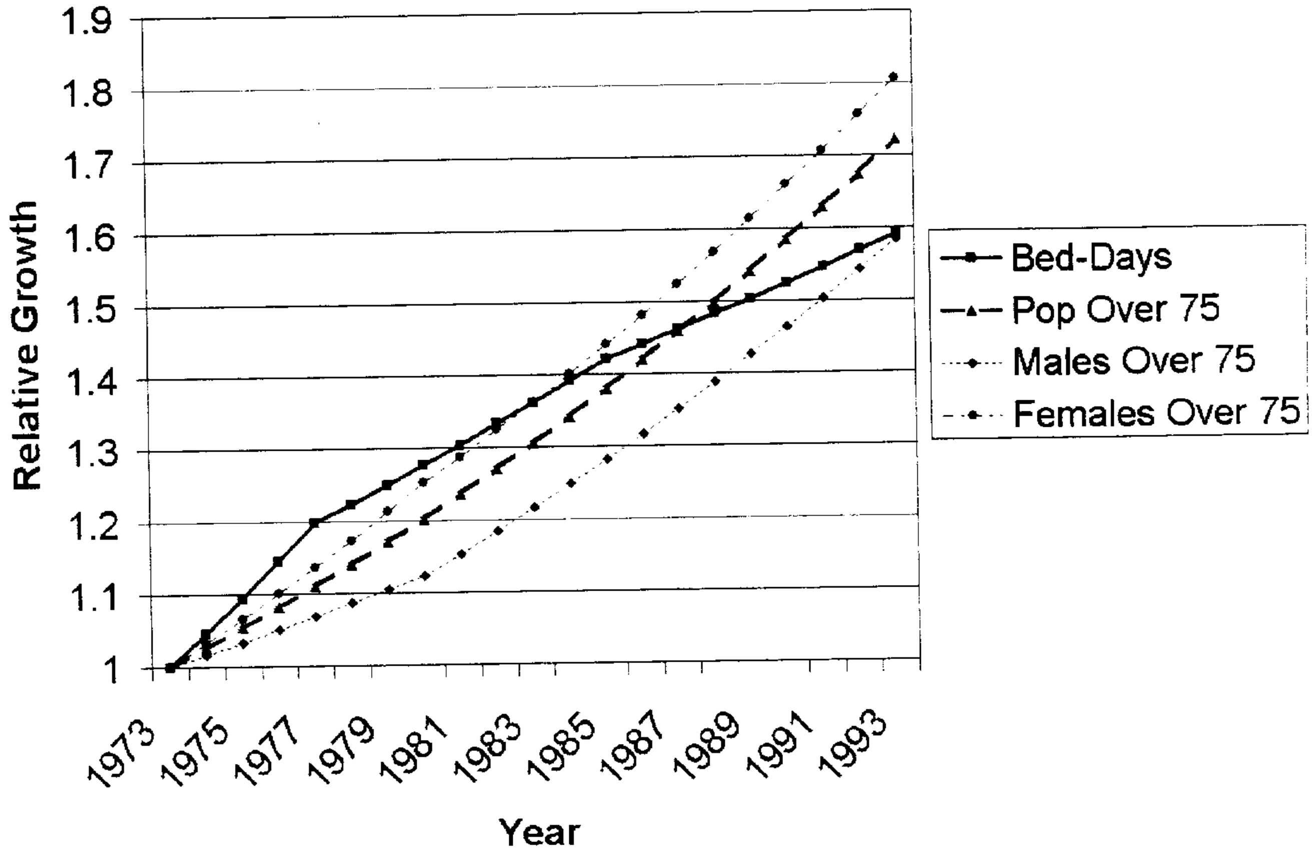

Figure 1: Relative Growth of Nationwide Nursing Home Bed-Days *versus* Relative Growth of Elderly Population (1973-1993).

bed-days in nursing homes to about 70 percent of bed-days by 1991.[4]

This paper provides a theoretical and empirical analysis of the relationship between aging and the growth of long-term care. The analysis will provide new insights into the forces which have contributed to the deceleration of nursing home output growth. The paper may be outlined as follows. Section 2 studies the impact of aging on the equilibrium in the long-term care market. We depart from the standard arguments that aging raises only the demand for care and show that aging directly increases the *supply* of home care, a primary substitute for market care, and thus reduces the demand for market care. This direct effect may be reinforced or counteracted by an indirect demand effect which depends on the covariance between longevity and frailty: if increases in longevity raise frail life-span more than healthy life-span, they will raise demand for market care, but if not, they will actually lower the demand for market care and reinforce the direct effect.

Section 3 generalizes the single-sex analysis of section 2 in order to consider the impact of aging on the supply of spousal care, which frail individuals may use as a substitute for market-based care. The theory here predicts that market-based output *contracts* with the longevity of the scarcer sex, typically the males, and that it expands with the *healthy* life-expectancies of the abundant sex, typically the females. This obtains because 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. By decreasing the differential between male and female life-spans, growth in elderly males reduces the need 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 raises demand for market care.

Section 4 provides empirical evidence from two independent data sources on the predicted impact of aging on equilibrium output growth. First, we consider the relation between aging and market output in US states from 1989-94 using aggregate state data from HCIA. We supplement this with an analysis of aging and long-term care within all US counties using data from the Bureau of Health Professions Area Resource File. Consistent with our predicted equilibrium effects of aging, we find on both the state and county-levels that a gender-neutral increase in the stock of old by one percent raises output by about 0.8 or 0.9 percent in the short-run, but only by 0.6 or 0.7 percent in the long-run. Furthermore, gender-specific growth in the old affects output in the predicted manner: the effect of female growth is positive but the effect of male growth is *negative*. Indeed, growth in the stock of old males often reduces market output more than one-for-one in many counties during this period. We find our estimated gender effects to be quite robust to various controls, including differences in frailty across states, and interstate differences in legislative barriers to entry in the nursing home industry. Given these estimated gender effects, the rapid deceleration in nursing home bed-days growth is consistent with the closing of the gender gap in growth rates. As illustrated in Figure 1, during the 1970s, the more abundant gender (females) grew much faster than the less abundant gender; this necessitated a more than proportionate expansion in market care. However, as male growth rates roughly reached female growth rates in the 1980s,[5] the relative supply of home care rapidly increased and caused a deceleration in the quantity of market-based care supplied.

The paper relates to an existing but rather scarce economic literature on long-term care.[6] 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.

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

[5] Initial evidence suggests the continued strengthening of this trend through the 1990s. Male growth among the elderly slightly exceeds female growth for the early '90s. (See Table 2.)

[6] See, for instance, Norton (1997), Wise (1992), Posner (1995), Garber (1994), Norton and Newhouse (1994), Gertler (1989), Gertler and Waldman (1992), and Weisbrod (1988).

Understanding the macro-level interactions between biological and economic forces in this market seems important for understanding the movement of output over time.

2 The Biology and Economics of Long-Term Care

While economists often claim that aging will raise the demand for market care, in this section we claim that the direct effect of aging is to *reduce* that demand. This direct effect may be reinforced or counteracted by an indirect demand effect which partly depends on the covariance between longevity and frailty.

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 $\nu \equiv E[S]$; and $T-S$, consequently, is the healthy years, with $\varphi \equiv \mu - \nu$ being the healthy life-expectancy. Let (H_t, F_t) denotes the stocks of healthy and frail individuals at time t. Under time-independent exit rates, these stocks change over time according to

$$\begin{aligned} \dot{H}_t &= b - \frac{1}{\varphi} H_t \\ \dot{F}_t &= \frac{1}{\varphi} H_t - \frac{1}{\nu} F_t, \end{aligned}$$

The variable b represents the size of the entering cohort,[7] and the inverse of healthy and frail life-expectancies define hazard rates into frailty (conditional on health) and mortality (conditional on frailty), 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. These steady-state effects generalize easily when the exit rate into disability or mortality is age-dependent. Assume 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 $S_T(t)$ and the bottom survival is the healthy lifetime survival $S_{T-S}(t)$.[8] The frail members of a cohort have completed their healthy durations, $T - S \geq t$, 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 healthy. Therefore, this stock is given by the shaded area between the two curves in Figure 2 and can be expressed as:

$$F = b \int [S_T(t) - S_{T-S}(t)]dt = b(\mu - \varphi) = b\nu$$

Similarly, since the area below the lower curve in Figure 2 represents the stock of healthy individuals, this stock must satisfy:

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

[7] We abstract here from the fertility effects of reductions in mortality; in other words, b is not a function of μ.

[8] $S_{T-S}(t)$ gives the *unconditional* probability of being healthy at time t.

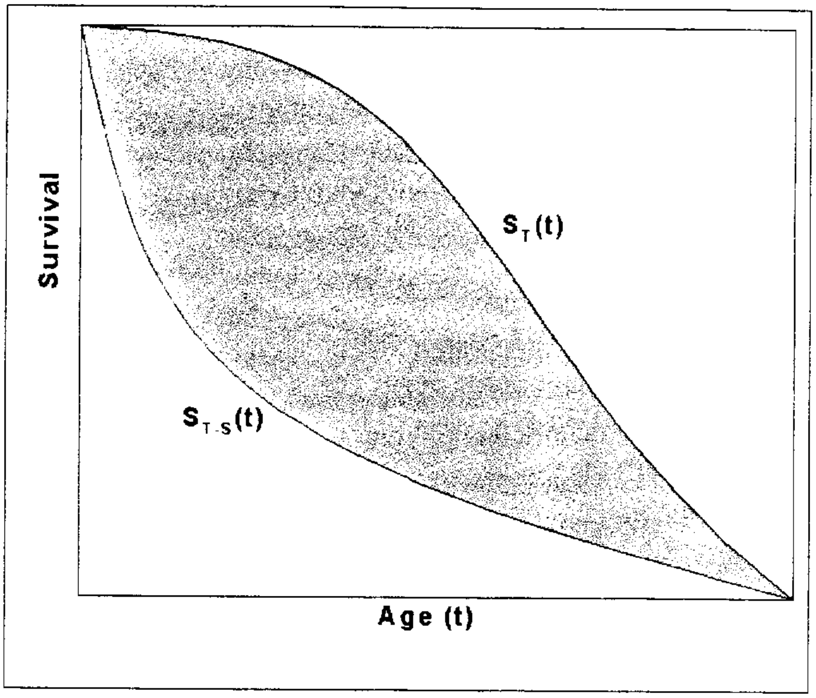

Figure 2: Overall and Healthy Lifetime Survival Functions.

The total demand for long-term care by frail individuals is defined by the stock of frail multiplied by their per-capita demand $d(p, H)$ which is assumed to fall in both price and the availability of substitute home care, which rises in the number of healthy individuals H: $d_p \leq 0$ and $d_H \leq 0$. The supply of care is represented by an upward sloping supply curve $Z(p)$ that comes from younger individuals outside of our model of the elderly population. The upward sloping supply curve reflects the labor-intensity of long-term care, in which labor costs represent almost all production costs. The equilibrium quantity and the price thus satisfy:

$$Fd(p, H) = Z(p) \tag{1}$$

Let $P(\mu, \nu)$ and $Y(\mu, \nu)$ denote the equilibrium price and quantity. Holding frail life-span ν constant, demand $Fd(p, H)$, and thus price and quantity, fall in longevity:

$$P_\mu \leq 0 \quad \& \quad Y_\mu \leq 0$$

This direct negative implication differs from the usual argument that aging *per se* raises demand and consequently, the equilibrium quantity and price. In fact, aging increases the stock of healthy caretakers relative to the stock of frail individuals who need care. In Figure 2, a rise in longevity pushes the lifetime survival curve S_T outward, holding the shaded area, which determines the demand for care, constant. Since healthy life-span rises and frail life-span remains fixed, the supply of caretakers rises relative to the quantity of frail individuals. This represents the *direct* effect of aging.

Changes in the size of this shaded area as the top survival curve S_T expands to the northeast represent the *indirect* demand effect of expanded longevity, the effect which we will now examine. The *dependence* between frailty and longevity determines the nature of this effect.[9] One may characterize this dependence through two conditional mean functions $\nu(\mu)$ and $\varphi(\mu) = \mu - \nu(\mu)$ which relate the average length of disabled and healthy lifetimes to life-expectancy. These functions reveal whether or not longer-lived cohorts are disabled longer. Since healthy and disabled life make up total life time, as in $\nu(\mu) + \varphi(\mu) = \mu$, the impact of longevity on the two must equal the longevity effect itself. Writing the equilibrium price and quantity as a function of longevity alone we get $P(\mu) \equiv P(\mu, \nu(\mu))$ and $Y(\mu) \equiv Y(\mu, \nu(\mu))$. When frailty is a function of longevity, the equilibrium price responds positively to longevity increases whenever frailty growth outweighs the growth of healthy home producers. Specifically, if we define the elasticity of demand with respect to healthy stocks as $\theta_H \equiv (-d_H)\frac{H}{d}$, and the elasticity of health stocks with respect to longevity as $\eta_H \equiv \left(\frac{\partial H}{\partial \mu}\frac{\mu}{H}\right)$, we can use the market-clearing condition in 1 to show that $\eta_P \equiv (\frac{dP}{d\mu})(\frac{\mu}{P}) \geq 0$ if and only if

$$\eta_F \equiv (\frac{dF}{d\mu})(\frac{\mu}{F}) \geq -\eta_H\theta_H = -\frac{\mu(1-\nu'(\mu))}{\mu - \nu}\theta_H$$

If we observe that the elasticity of frailty with respect to longevity η_F equals $\eta_\nu \equiv \nu'(\mu)\frac{\mu}{\nu}$, the elasticity of expected frail lifetime ν, the expression simplifies to:

$$\eta_\nu \geq -\frac{\mu\theta_H}{\mu - \nu + \nu\theta_H} \tag{2}$$

The right-hand side is the elasticity of frailty with respect to longevity and thus represents the percentage change in the pool of frail individuals, the shaded area in Figure 2, induced when longevity rises by one percent. The left-hand side represents the percentage decrease in demand induced by any gains in healthy lifetime resulting from the increase in longevity. This shows more precisely the direct and indirect effects of aging on demand, and hence price. Price rises if the indirect positive effect of longevity on the demand side, η_F , exceeds the direct negative demand effect due to the growth in substitute home production. If a one percent increase in healthy stocks reduces market demand by one percent, so that $\theta_H = 1$, we see from 2 that in a steady-state where frailty is held constant, a one percent increase in life expectancy increases healthy stocks just enough to reduce demand by one percent. In this case, longevity gains increase the price of long-term care if a one percent increase in longevity increases frail lifetimes by more than one percent, or $\eta_\nu \geq 1$.

If we differentiate the market demand function $Fd(p, H)$ with respect to longevity, and substitute the equilibrium elasticities of demand into the resulting expression, we can show that the elasticity of equilibrium output to longevity is given by:

$$\eta_Y \equiv (\frac{dY}{d\mu})\frac{\mu}{Y} = \theta_H\eta_H + \theta_p\eta_P + \eta_F$$

$\theta_p \equiv d_p p/d$ represents the price-elasticity of per-capita demand and $\theta_H\eta_H$ is the direct elasticity of demand with respect to increases in health. The first effect is the *direct* effect of longevity on the demand for care. This direct effect always decreases demand and thus quantity. The second term is

[9] In the conclusion, we briefly discuss predictions for this dependence when increasing life-spans induce larger health investments and hence lower disability. Recently, this type of dependence has been documented across US cohorts by Manton, *et al.* (1997). If the covariance is expressed as a linear conditional mean function, then $E[S|T = t] = \beta_0 + \beta_1 t$ and hence $\nu(\mu) = \beta_0 + \beta_1\mu$. These parameters may then be estimated by regressing length of disability on age of death (see Philipson, *et al.* (1997)).

the *indirect* effect of aging on demand by means of its effect on price. If longevity raises healthy home production more than it expands frailty, the price falls and the price effect on quantity is positive. The opposite occurs if longevity expands frailty more than healthy lifetimes.

To illustrate, consider the case in which the stock of frail individuals rises proportionately with increased longevity, $\eta_F = 1$, and in which demand is completely inelastic, $\theta_p = \theta_H = 0$. In this case, there is no price response, as the rise in home production by healthy individuals balances the rise in demand by frail. Consequently, since per-capita demand is inelastic a one percent increase in longevity induces a one percent increase in the equilibrium quantity. These assumptions seem to be the implicit market assumptions made by theorists who project growth in the quantity of long-term care strictly from population growth.

3 Marriage Markets & Long-term Care Markets

This section extends the single-gender analysis of Section 2 by studying the impact of marriage markets on the relation between long-term care markets and aging. Due to higher female longevity and the matching of older males with younger females, there will be more female care-givers for males than male care-givers for females. Market production makes up for this relative scarcity of male care-givers in home production. This implies three main predictions about the impacts of increased longevity on the total steady-state output. First, when female healthy life-span rises, market output of long-term care rises, but when male healthy life-span rises, market output of long-term care actually *falls*. Second, gender-neutral growth in longevity results in (weakly) less than proportionate growth in the output of long-term care. Third, the relative scarcity of men in home production causes the proportion of women in nursing homes to exceed the proportion of women among the frail.

To examine the impact of marriage matching on output, we will consider an elderly population in which all healthy men and women are matched until the supply of the smaller gender is exhausted. As before, there are three states of health for every individual: healthy, frail, and deceased. Generalizing the previous single sex case, the distributions of overall and frail lifetimes (T_m, S_m) and (T_f, S_f) for males and females can be characterized by their means (μ_m, ν_m) and (μ_f, ν_f).[10] As before, the relevant question is of the direct and indirect effects of a rise in female and male longevity, μ_f and μ_m,on market output.

Healthy couples, not just healthy individuals, are assumed to enter this population every period. If one partner falls into frailty before the other, we assume that the healthy partner cares for the frail one at home. If both partners become frail, we assume that both require market-based long-term care: for the purposes of our analysis, we can consider such partners as unattached. Finally, every period, all healthy widows are matched to healthy widowers until the supply of healthy widowers runs out: we assume that the number of healthy widows always exceeds that of healthy widowers. Let b now represent the size of the entering cohort of *couples*, rather than individuals. Let C be the stock of healthy couples; let C_f be the stock of couples where the female is disabled, and let C_m be the stock of couples where the male is disabled. The dynamics of these couples under time-independent transitions

[10] Conditional on reaching 65, the healthy life-expectancy at 65 is often estimated to be about two-thirds of remaining years, so that one-third of life after retirement is disabled. Conditional on this age, males often have higher *proportions* of healthy life than females. In our notation, the U.S. numbers for 1994 were about $\varphi_m = 9$ and $\nu_m = 6$, for males and $\varphi_f = 11$and $\nu_f = 7$ for females. Therefore, $\mu_m = 15$ for males, and $\mu_f = 18$ for females. By implication, after having reached age 65, US men and women spend about 60 percent of their remaining lives in the healthy state, although females are disabled longer in absolute time than are males (United Nations (1995)).

is then given by:

$$\dot{C} = b - (\varphi_f^{-1} + \varphi_m^{-1})C + \nu_f^{-1}C_f \tag{3}$$

$$\dot{C}_f = \varphi_f^{-1}C - \nu_f^{-1}C_f - \varphi_m^{-1}C_f \tag{4}$$

$$\dot{C}_m = \varphi_m^{-1}C - \nu_m^{-1}C_m - \varphi_f^{-1}C_m \tag{5}$$

Observe that healthy males whose frail wives die, $\nu_f^{-1}C_f$, are always matched again, because we assume that there are enough healthy widows to marry all available healthy widowers; healthy widows who cannot be matched will enter the pool of single, healthy women, to be described below.

Denote by H_m and H_f the stocks of unattached healthy men and women, respectively. Also, denote by F_m and F_f the stocks of unattached frail men and women, respectively. Only the unattached frail men and women enter market care, so the output of market care is given by: $F_m + F_f$. The transitions for single, healthy men and women will be given by:

$$\dot{H}_m = -\varphi_m^{-1}H_m \tag{6}$$

$$\dot{H}_f = \nu_m^{-1}C_m - \nu_f^{-1}C_f - \varphi_f^{-1}H_f \tag{7}$$

Since there are always enough healthy widows for healthy widowers to get matched again, there is no source for healthy widowers in this model. In a steady-state, therefore, there will not be any healthy unattached widowers. Observe that healthy widows unable to find a mate enter the pool of healthy unattached females.

Finally, we have the stocks most important for the determination of market care: the stocks of frail, unattached men and women. These stocks change over time according to:

$$\dot{F}_f = \varphi_f^{-1}H_f + (\varphi_m^{-1}C_f + \varphi_f^{-1}C_m) - \nu_f^{-1}F_f \tag{8}$$

$$\dot{F}_m = \varphi_m^{-1}H_m + (\varphi_m^{-1}C_f + \varphi_f^{-1}C_m) - \nu_m^{-1}F_m \tag{9}$$

These transitions formally illustrate the three paths to unattached frailty for any man or woman: an agent can be unattached and healthy, but then fall into frailty; she can be healthy taking care of a frail spouse, but then fall into disability; she can be healthy, but taken care of by a healthy spouse, who may himself fall into frailty.

As was the case for the single-sex analysis, market output is the intersection of the demand curve of frail individuals with the supply curve of younger individuals

$$(C_f + C_m)d^C(p) + (F_f + F_m)d(p) = S(p)$$

Now the per capita demand for long-term care is represented by d^C for a frail individual with a spouse and d for a single frail individual. The number of healthy old helps to lower the per capita market demand provided that for all prices, $d^C(p) < d(p)$. To make things simple in this section, we will assume that $d^C(p) = 0$, so that frail individuals with a healthy spouse never enter nursing

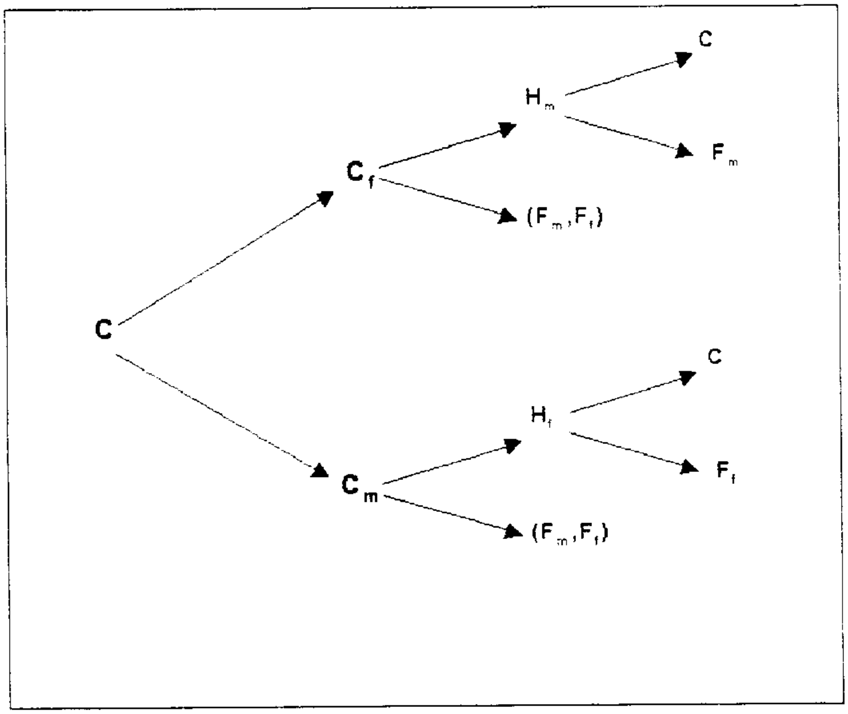

Figure 3: Transitions between Health States for Couples.

homes. Moreover, in doing the formal analysis below, we will assume that every frail single has the same demand for care $d(p)$. As a result, a given percentage change in frail singles must induce the same percentage change in total market demand. In other words, denoting the equilibrium price and quantity by $P(\mu_m, \mu_f)$ and $Y(\mu_m, \mu_f)$, we need only characterize the movement of $F_f + F_m$ in order to characterize the movement of market quantity and price: if $F_f + F_m$ rises, then total demand must rise, so that market price and quantity rise; if $F_f + F_m$ falls, total demand falls, so market price and quantity must fall.

The relations between the various states are illustrated in Figure 3 below, which shows how healthy couples move into the care-giving, -receiving state, and from here move into the various unattached states. In this model, only women enter the unattached healthy state. This feature drives the first prediction: *increases in male longevity lower market demand, but increases in female longevity raise it.* This will hold true provided that disabled unattached males make up an empirically insignificant proportion of the total population of disabled and unattached people. Suppose that male healthy life-span increases. This increases the supply of male care-givers for disabled females and thus directly reduces the stock of females in market-based long-term care. Of course, the increase in male healthy life-span reduces the probability of frailty, but it also raises the steady-state stock of males, so the impact on the number of frail males is ambiguous. Using our assumption that women comprise the vast majority of unattached frail people, the increase in unattached frail males will not offset the reduction in unattached frail females. By similar reasoning, when female healthy life-span increases, the scarcity of male care-givers becomes exacerbated. More women, left without healthy mates, will have to enter market-based long-term care when they become disabled, because they will have outlasted their spouse's healthy lifetime. Market-based care increases, again contingent on the relative insignificance of disabled, unattached males: an increase in female healthy life-span reduces the stock of such males, but this decrease will be overwhelmed by the increase in unattached frail females.

To show this more precisely, we outline the formal proofs, details of which may be found in the Appendix. Market-based care occurs only for unattached frail men and women. These steady-state stocks must satisfy:

$$F_m = \nu_m(\varphi_m^{-1} C_f + \varphi_f^{-1} C_m)$$

$$F_f = \varphi_f^{-1} H_f + \frac{\nu_f}{\nu_m} F_m$$

Now, market care demand is driven by the sum of unattached frail males and females. We assume throughout that the stock of unattached frail females greatly exceeds the stock of unattached frail males. Therefore, assuming ν_f and ν_m are roughly equal, the term $\varphi_f^{-1} H_f$ will dictate the movement of market care. In words, the number of healthy widows entering disability (which is exactly $\varphi_f^{-1} H_f$) governs the movement of market care, provided that the number of unattached frail males is relatively small. Moreover, the number of healthy widows falling into frailty rises in female healthy life-span and falls in male healthy life-span. Formally, we show the following proposition in the appendix:

Proposition 1 In a steady state with relatively few frail widowers, increases in male longevity reduce market output and price, but increases in female longevity raise market output and price: $\frac{\partial Y}{\partial \mu_m}, \frac{\partial P}{\partial \mu_m} < 0$ and $\frac{dY}{d\mu_f}, \frac{\partial P}{\partial \mu_f} > 0$.

In words, as women become healthier, there will tend to be more healthy widows, and in spite of the fall in the probability of becoming frail, an increase in healthy life-span raises the number of healthy widows entering frailty. As men become healthier, on the other hand, there will be fewer healthy widows, and thus fewer healthy widows entering frailty. Increases in female health raise the demand for male care-givers, exacerbate the scarcity of these care-givers, and thus result in increased demand for market care. In contrast, increases in male health raise the supply of male care-givers, reduce the scarcity of these care-givers, and thus result in *less* demand for market care provision.

The second prediction states that, provided that the ratio between healthy and frail lifetimes remain unaffected, *gender-neutral population increases have proportionate effects on market demand, and thus weakly less than proportionate effects on the output of care.* The exactly proportionate effect will obtain provided that demand among single frail people is inelastic. To understand the impact of both a gender-neutral population increase, and a relative increase in the female population, first note the following useful facts, shown below (see note 11): in a steady-state, the sex ratio $\frac{females}{males}$ is given by $\frac{\varphi_f + \nu_f}{\varphi_m + \nu_m}$, while the frailty ratio for either sex, $\frac{frail}{healthy}$, is given by $\frac{\nu}{\varphi}$; finally, the steady-state stock of either gender is given by $b(\varphi + \nu)$. A population increase which is gender-neutral and frailty-neutral must affect neither the sex ratio nor the frailty ratio. Such an increase may be represented either as a fixed percentage increase in φ_f, φ_m, ν_f, and ν_m, or as a fixed percentage increase in b. The first part of our prediction may now be stated as the following proposition:

Proposition 2 *Frailty-neutral and gender-neutral aging implies proportionate growth in frail singles and total market demand, and thus (weakly) less than proportionate growth in market output.* $\frac{dY}{db}(\frac{b}{Y}) \leq 1$

This proposition obtains, because increasing longevity without affecting either the sex ratio or the frailty ratio merely increases the existing population by a certain percentage and increases the existing population of frail singles by the same percentage. Therefore, total market demand rises proportionally, and depending on the elasticity of demand, total market output rises weakly less than proportionally

We can illustrate this proposition with a simple example, which demonstrates that frailty-neutral and gender-neutral growth results in exactly proportionate growth in frail singles. Suppose we begin with 100 women, 90 men, and a female frailty ratio of $\frac{1}{9}$ (this of course translates into a 10% frailty proportion). For simplicity, suppose there are no unattached frail males. Therefore, we begin with 10 single women, one of whom is frail. Total market demand is just $d(p)$, the demand of the one frail single woman. Now suppose that φ_f, φ_m, ν_f, and ν_m all double. Since the sex ratio and the 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, 2 of the 20 unattached women will be frail, so market demand moves to $2d(p)$. Since every frail single has the same per capita demand for long-term care, market demand exactly doubles. Observe also that if the frailty ratio falls with the increase in population, total demand increases less than proportionately; conversely, if the frailty ratio rises with the increase in population, total demand increases more than proportionately. In other words, the frailty ratio governs the magnitude of the effect induced by a gender-neutral rise in population. Now, observe that a proportionate change in total demand results in a weakly less than proportionate change in market output, where the exactly proportionate case obtains if demand is inelastic.

Many times longevity increases are not in fact gender-neutral. In the Appendix, we outline and prove Proposition 3, which argues that *relative increases in female longevity increase total demand more than proportionately.* Such an increase, holding the female disability ratio constant, may be represented as a fixed percentage increase in φ_f, and ν_f, while φ_m, and ν_m remain constant. This result follows from the scarcity of males. If the stock of females increases, *holding males constant*, none of the new females have healthy spouses. Therefore, the rate of entrance into the frail single state will be equal to the frailty rate itself. Since at least some frail females in the existing population are cared for by healthy males, the rate of single frailty in the existing population must be strictly less than the frailty rate. Since the new females are more likely to enter single frailty than the existing females, the population of frail singles must increase more than proportionately, and thus total market demand increases more than proportionately. Observe that here we once again need the assumption that there are few frail unattached males, because the effect on such males of an increase in female longevity may be negative.

Again, consider an illustrative example. Suppose we have 100 females, 90 males, and a female frailty ratio of $\frac{1}{9}$. Of the 10 unattached females, only one is frail. Suppose φ_f and ν_f both increase by 10%, so the frailty ratio is unaffected, but we have a 10% increase in the steady-state stock of females. Therefore, we have 10 additional females, all of whom are unattached. Two of the 20 total unattached females are frail, so we have a single frail population of 2. Observe that a 10% increase in female stock resulted in a 100% increase in the stock of frail singles. As long as some women are married, the effect must be more than proportional, but the size of the effect depends on how many women are married. In the extreme case of no married women, the effect is exactly proportional, because the rate of single frailty rate in the population is just the frailty rate itself. In sum, the more healthy men available, the greater will be the effect on nursing home residents of a given increase in females.

The last implication of interest concerns how aging impacts the gender composition of market care.

We find that *the proportion of women in nursing homes weakly exceeds the proportion of women among the frail.* Since women will be unattached more often than men, the proportion of women among the single frail must exceed the proportion of women among the frail. Observe that entrance into single frailty requires both frailty and the absence of a healthy spouse. If there exists a scarcity of males, women will be less likely to be married. As a result, even controlling for differences in frailty, women will be more likely to enter single frailty, and thus more likely to receive market-based long-term care. At the extreme, where we have no unmarried healthy people, both genders are equally likely to be married, so that only the incidence of frailty determines the probability of entering single frailty. In this case, the proportion of women in nursing homes must be the same as the proportion of women among the disabled, because both sexes have equal access to marriage.

It is simple and instructive to show this claim formally. In our model, the ratio of females to males among the total population, δ, among the frail, δ_F, and among those demanding market care (i.e., those who are single and frail), δ_M are given by

$$\delta \equiv \frac{C + C_m + C_f + F_f + H_f}{C + C_m + C_f + F_m + H_m}$$

$$\delta_F \equiv \frac{C_f + F_f}{C_m + F_m}$$

$$\delta_M \equiv \frac{F_f}{F_m}$$

Given our work in the previous section, it is not difficult to calculate these quantities explicitly. In a steady-state, the number of healthy individuals of sex i is given by $\varphi_i b$, while the number of disabled individuals of sex i is given by $\nu_i b$.[11]

Using these facts, and the steady-state values for F_f and F_m (see the appendix), we can now explicitly describe the ratios δ, δ_F, and δ_M:

$$\delta = \frac{\mu_f}{\mu_m}$$

$$\delta_F = \frac{\nu_f}{\nu_m}$$

$$\delta_M = \frac{\nu_f}{\nu_m} + \frac{\varphi_f^{-1} H_f}{F_m} = \delta_F + \frac{\frac{\varphi_f}{1+\frac{\nu_m}{\varphi_f}} - \frac{\varphi_m}{1+\frac{\nu_f}{\varphi_m}}}{\frac{\nu_m}{\nu_f^{-1}+\varphi_m^{-1}} + \frac{\nu_m}{\nu_m^{-1}+\varphi_f^{-1}}}$$

Observe that the first two expressions are independent of marriage: a society without marriage would display these patterns. However, the ratio of women to men demanding market care, δ_M, depends on the relative scarcity of mates in the marriage market. In particular, the more healthy widows H_f we see, the more the proportion of women demanding market care exceeds the proportion of

[11] To illustrate the reasoning behind these relations, we will prove them for men. The transition for healthy men is given by $\dot{C} + \dot{C}_f = b - \varphi_m^{-1}(C + C_f)$, which implies the steady-state value $C + C_f = \varphi_m b$. The transition for total men is given by $\dot{M} = b - \nu_m^{-1}(C_m + F_m)$. In a steady-state, this implies $C_m + F_m = b\nu_m$.

women among the disabled, and the more the proportion of women in nursing homes will exceed the proportion of women among the disabled. If both genders are perfectly matched, so no healthy people are unmarried, $\delta_M = \delta_F$ and the two ratios exactly coincide. However, any inequality in the marriage market between genders will cause δ_M to diverge from δ_F. Intuitively, marriage provides a type of "insurance" against the possibility of market-based long-term care: a married person will not enter market-based long-term care unless he *and* his spouse become frail. If H_f is large, the insurance is available to fewer women than men. As a result, the ratio of women to men in market-based long-term care facilities will weakly exceed the ratio of disabled women to disabled men.

The predominantly female demand for market care is most commonly explained as the result of lower female mortality, $\mu_f \geq \mu_m$. This argument misses a crucial part of the story. First, while the condition $\mu_F \geq \mu_M$ turns out to be both necessary and sufficient for there to be more females than males ($\delta \geq 1$), it is neither necessary nor sufficient for there to be more disabled females than males: this condition is governed entirely by the expected duration of frail lifetimes across gender. Precisely, we have $\delta_F > 1$ if and only if $\nu_f > \nu_m$, but in itself the relation between μ_M and μ_F does not affect δ_F. Second, even relative frail lifetimes do not by themselves account for the fact that the share of women in nursing homes exceeds the share of women *among the disabled*. The disadvantages of women in the marriage market cause women to be overrepresented in nursing homes, relative to their representation among the disabled.

4 Empirical Analysis

This section considers the impact of aging empirically. We consider both cross-sectional and longitudinal evidence of the effects of aging on the quantity of long-term care provided by considering the short-run growth of the old population within US states and the long-run growth of the old population within US counties.[12] Principally, we find that a one percent increase in the population of elderly men *reduces* long-term care output by approximately 1.3%, while a one percent increase in the population of elderly women *raises* long-term care output by approximately 1.2%.[13] Therefore, a "gender-neutral" population increase of one percent raises output by about 0.9%. When we control for Medicaid subsidization, the gender-neutral effect falls to around 0.8%. We also find that these effects are robust to controls for Medicaid subsidization, legislative entry barriers in long-term care, and variation in rates of frailty across states.

4.1 Aging and Long-Term Care Across States

The data we use for the analysis at the state level come from HCIA's *Guide to the Nursing Home Industry*, as well as electronically published supplements to that guide. The data set is in panel form, with values for every state and the District of Columbia over the years 1989 through 1994. The summary statistics for the data are reported in Table 1 below.

INSERT TABLE 1 HERE

The HCIA data set contains population breakdowns by gender and age for every state. We use these to construct statistics for male and female populations for three age groups: over 65; over 75;

[12] It serves to point out that the analysis here does not aim to identify unobservable supply and demand schedules but to estimate the impact that observable shifts of these schedules have on equilibrium output.

[13] The magnitudes sometimes differ across data sets and specifications, but these are "middle-of-the-road" estimates: the signs of the estimates, however, are invariant to data sets and specifications.

over 85. In addition, HCIA collects data from state governments on the total beds available statewide in the nursing home industry. We use these figures, multiplied by 365, to calculate bed days available yearly in each state. To calculate yearly bed days filled, we need information on capacity utilization. Unfortunately, no available statistic calculates this number for all nursing homes in each state. Instead, for each state, HCIA calculates the median occupancy rate in a sample of nursing homes, where the sample for a given year is the set of homes which submitted Medicare or Medicaid claims that year. We multiply this sample median by yearly bed days available to construct a measure of yearly bed days filled, reported in the first row of Table 1.

By aggregating the statewide data, we can also gain some initial insight into gender growth and nursing home growth for the early '90s. The nationally aggregated statistics are presented in Table 2.

INSERT TABLE 2 HERE

Over the early '90s, we see the continued balancing of gender growth, and the continued deceleration of nursing home growth which would be implied by this balancing. Table 2 demonstrates that male stocks actually grew faster than female stocks over this period; the closing of the gender gap in growth rates, which began in the '80s, has been completed to such an extent that we now witness the very beginnings of a closing in the gender gap in elderly *stocks*. Based on the theory in Section 3, this increasingly balanced gender growth leads us to predict the continuing deceleration of growth in nursing home bed-days. Indeed, table 2 also quantifies the extent to which nursing home bed-day growth continues to decelerate. The 5-year bed-day growth rate of 7.33% corresponds to an annual growth rate of about 1.42%, a far cry from the 4.61% annual growth of the '70s, and even the 2.17% annual growth of the '80s. The trends in the summary data appear to confirm our theory of gender effects.

We will primarily be interested in testing the hypothesis that the covariance between the male elderly population and yearly bed days will be negative, while the covariance between the female elderly population and yearly bed days will be positive. We will estimate the panel specification:

$$LogBed_{it} = \gamma_1 + \gamma_2(LogMale_{it}) + \gamma_3(LogFemale_{it}) + \varepsilon_{it} \tag{10}$$

The regression will be run across states and time, so i indexes the 51 states,[14] while t indexes the years of the data set, 1991 through 1994.[15] *LogBed* is the natural logarithm of yearly bed days produced, while *LogMale* and *LogFemale* refer to the natural logarithm of stock of male and female elderly populations, respectively. Results of these and other regressions appear in Table 3. The results of this table support our predictions that increases in elderly males do in fact decrease the output of long-term care, while increases in elderly females raise it.

INSERT TABLE 3 HERE

For all three age groups, the point estimates have different signs: γ_2 comes in as negative, while γ_3 comes in as positive. Once we note that $\gamma_2 + \gamma_3$ delivers the elasticity of nursing home bed-days with respect to a gender-neutral increase in population, we can say that a gender-neutral 10% growth in the population over 65 results approximately in a 9.2% increase in bed-days; however, the 10% increase in males results in a 19.3% *decrease* in bed-days, while the 10% increase in females results in

[14] We include as a "state" the District of Columbia

[15] In the HCIA data set, values for total statewide nursing home beds begin in 1991.

a 28.5% *increase*. For the over 75 age group, 10% gender-neutral growth results in a 9.3% increase in bed-days, of which 21% owes itself to the growth in females, while -11.7% owes itself to the growth in males. For the over 85 age group, 10% gender-neutral growth creates a 9.6% increase in nursing home bed-days, of which -5.7% is due to male growth, and 15.3% is due to female growth. Evidently, the negative effect of male growth becomes attenuated when we consider older age groups, perhaps because growth at these ages represents growth in individuals who are more frail, and thus less able to provide home care.

We should now mention that while the first-difference regressions tend to deliver point estimates with negative signs for males and positive signs for females, they do not come out as statistically significant, probably due to the small sample size, and due to a lack of variation within states over this short period of time. Due to the short time-span of our data, most of the explanatory power in our regressions thus comes from cross-state variation in levels. This relative absence of time-series variation leads us to consider a longer time span in the county data analyzed below.

4.2 Long-Run Growth in Aging and Long-Term Care in Counties

The data we use for the county-level analysis come from the Bureau of Health Professions (United States Department of Health and Human Services, Health Resources and Services Administration) Area Resource File, 1940-1995 edition. From this file, we have taken county-level data on the long-term care output and demographic characteristics of each county. The file contains county-level data on the number of long-term care facilities, the total number of beds in all such facilities, and the total number of residents in all such facilities. A rough occupancy rate may be calculated by dividing residents into beds. Since these rates are very close to one hundred percent, we take total beds to be total beds occupied, and use this series to calculate total yearly bed days. 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 also those 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 linearly interpolated by county to construct series for 1971, 1973, 1976, 1978, 1980, 1982, 1986, and 1991. These data are summarized in Table 4.

INSERT TABLE 4 HERE

Using the county as the unit of the analysis, we will first reexamine the regression model given in equation 10. We will run this regression using the data on males and females over the ages of 65 and 75 (in this data set, we do not have data by gender on the population over age 85 for all the years in question), using all counties and all years for which we have long-term care data. In this section, since we have county data, we will also be able to control for state fixed effects. The wide variation across states in nursing home regulation suggests that ε_{it} contains a state fixed effect; that is, $\varepsilon_{it} = \mu_i + \nu_{it}$. We eliminate the fixed effect μ_i by running a within-state specification. The results are presented in Table 5. Basically, this table demonstrates that the panel effects at the county level complement and greatly strengthen our earlier state level findings. Increases in males reduce the output of long-term care while increases in females raise it. Furthermore, it offers evidence in support of our prediction that gender-neutral increases in population result in less than proportionate increases in long-term care, while female increases result in more than proportionate increases.

INSERT TABLE 5 HERE

The first two columns of Table 5 show us that a gender-neutral 10% increase in population over the ages of 65 and 75 raises long-term care bed-days by about 9%. This within-state estimate closely follows the state-level estimates from the previous section. The qualitative findings of section 4.1 are further confirmed by these data.[16] A 10% increase in the stock of males over 65 *reduces* bed-days by 13.3%, while a 10% increase in the stock of females over 65 increases bed-days by almost 22.5%. As before, we find that the negative effect exerted by males falls, by more than 60% here, for the older over 75 age group; since increases in older age groups would tend to reflect increases in frail life-span more than increases in healthy life-span, this tendency makes sense in the framework of our model. Moreover, we find that for older women, the effect of growth in females moves closer to proportionality, as we predicted. The increasing scarcity of healthy men at the older ages moves the female effect toward proportionality, as fewer and fewer women have healthy mates.

Since we are concerned with explaining a fundamental shift in the relation between output growth and population growth over time, and since our variables are clearly not stationary, we should check if our estimates are affected by this non-stationarity. The results in Table 5 demonstrate that the magnitudes of our coefficients are quite stable and appear not to be affected by detrending. In the third and fourth columns, we add a linear time term to the regression: effectively, this constructs for every variable a trend series with a constant twenty-year growth rate, and then removes this trend series from the dependent and independent variables. In the fifth and sixth columns, we run regressions within states *and* years; this operation detrends the dependent and independent variables by computing and removing a single constant within-state growth rate for each year. The gender-neutral population elasticity is not changed by either of the two detrending operations, and the gender-specific elasticities do not change by more than 0.1. Moreover, from the third and fourth columns of Table 5, reporting regressions which control for a constant exponential time trend, we find that after accounting for the impacts of gender growth, little time-series variation remains in the nursing home output series. Observe first that the R^2 values do not change with the introduction of detrending. Moreover, we find that controlling for gender growth over age 75, the level of output in 1991 is just 4% below its predicted level. Once we control for gender growth over age 65, we find that nursing home output in 1991 appears to be 8% *above* where it should be. There does not appear to be economically significant non-stationarity in the error term. The residual is not correlated with time strongly enough to compromise our estimates.

In sum, the estimates in Table 5 demonstrate that the change in gender growth can account for a significant portion of the deceleration in nursing home bed growth. Consider the within-state estimates for the regression on males and females over 75. This actually delivers a fairly conservative estimate of the gender effects: a male elasticity of -.48 and a female elasticity of 1.44. As Figure 1 illustrates, the population of males over 75 grew by 18.09% during the '70s, while the population of females grew by 37.85%, and nursing home bed-days grew by 45.98%; during the '80s, males grew by 30.30%, while

[16] Within-county estimates, also reported in table 5, are qualitatively similarly. Of course, the within-county estimates rely exclusively on variation across time, rather than the more reliable variation within states at a single point in time. Therefore, we regard as more reliable the within-state analysis, to which we limit our discussion in the text.

females grew by 32.57%, and nursing home bed-days grew by 19%. The elasticities along with the observed growth in gender stock imply a predicted growth in nursing home beds of 45.86% in the '70s, and 32.36% in the '80s. The model predicts a slowdown in growth of 13.5%. This represents over half of the actual 26.98% decline. Evidently, most of the slowdown may be explained in this extremely simple setting using conservative gender elasticities. Clearly, gender elasticities of larger absolute value will predict greater slowdowns.

Recall from the previous section that we did not find significant correlation between gender stocks and long-term care when we examined the data in differences. In that section, we only had access to data over a 5-year period. In this section, we repeat the analysis using 20-year differences at the county-level. That is, we have observations on over 3000 counties, and each represents a difference from 1971 to 1991. The results of these regressions are reported in Table 6. This table confirms, for the long-run, our predictions about the negative effect of male growth and the positive effect of female growth, as well as our predictions about the less than proportional effects of gender-neutral population growth and the disproportional effects of female population growth.

INSERT TABLE 6 HERE

Again, growth in male stocks is associated with decline in long-term care, and growth in female stocks is associated with growth in long-term care. For the simple first-difference case, the coefficients are highly significant for both age groups, with R^2 of near 0.8 for both regressions. The insignificance of the previous section probably owes itself to the long-term nature of these processes: growth in gender stocks probably do not result in short-term changes in long-term care. At this longer 20-year window, they have a significant impact. One puzzle remains, however. While the results for absolute differences seem convincing, and while the coefficients and significance levels are qualitatively similar for the regressions in log-differences (growth rates), the R^2 for these log-difference regressions drop dramatically, to around 0.11. This makes some sense when we observe that over 15% of the sample is lost, because the number of bed days in 1971 is zero for about 440 counties in the data set. In addition, and probably more importantly, since some counties have very small values for elderly people or bed days in 1971, the logarithmic form introduces a great deal of instability. In spite of the lowered R^2, the qualitative results support the earlier analysis. We find that a gender-neutral 10% rise in the stock of people over the age of 65 is associated with an 8% rise in the growth rate of long-term care, lower than the 9% short-run effect. Over a longer time-horizon, we expect the elasticity of demand for nursing home use to be lower, since substitutes such as family care may not be feasible in the short-run. Moreover, in the long-run, healthy widows and widowers have a chance to find new healthy mates. Of this total 8% effect, a 10 percentage point increase in the growth rate of the male stock decreases the growth rate of long-term care by 4%, while a similar increase in female growth raises the growth rate of long-term care by 12%. The effects are smaller for the over 75 age group, probably because growth in this age group represents growth in the frail population rather than the healthy population.

4.3 Increased Medicaid Subsidization and Output Growth

We have yet to investigate the dramatic increase in Medicaid subsidization over this 20-year period. To gain an initial understanding of the nationwide magnitudes involved, examine Table 7, which reports various quantities aggregated from the county-level to the national level.[17]

[17] The 1971 Medicaid data are drawn from NCSS (1974), and the 1991 Medicaid data are drawn from HCIA (1996). The intermediate years are linearly interpolated.

INSERT TABLE 7 HERE

Observe the dramatic increase in Medicaid subsidization over this 20 year window: Medicaid bed-days have more than tripled, while the share of bed-days subsidized by Medicaid has nearly tripled. Given this massive growth in subsidization, it is surprising that nursing home bed-days have not exploded in growth. Compared to the population over age 75, which represents the largest group of long-term care consumers, nursing home bed-days have actually increased less than proportionally, in spite of these massive increases in subsidization. Even if we consider the population over age 65, nursing home bed-days have only grown about 10% percent more over this 20-year time span.

We can gain a more precise understanding of the increase in bed-days accounted for by gender growth and Medicaid growth by explicitly examining the impact of gender growth on nursing home bed-days controlling for Medicaid, and *vice-versa*. In other words, we run the following regression:

$$LogBed_{it} = \gamma_1 + \gamma_2(LogMale_{it}) + \gamma_3(LogFemale_{it}) + \gamma_4(MedShr_{it}) + \varepsilon_{it} \tag{11}$$

Here, the variable $MedShr_{it}$ represents the year t statewide share of total nursing home bed-days paid for by Medicaid, in the state of county i. We will again examine this specification within states in order to control for state fixed-effects. The results of these regressions are reported in Table 8. Primarily, we find that Medicaid subsidization exerts a comparatively small effect on nursing home growth, once we control for gender effects. In fact, gender effects dominate Medicaid effects over this 20-year period.

INSERT TABLE 8 HERE

If we compare the within-state gender elasticities in Table 8 to those in Table 5, we find that the gender-neutral elasticities for both age groups are almost identical. The gender-specific elasticities in Table 8also remain within the range of the estimates given in 5. For those over 65, 10% growth in males decreases output by 12.5%, while 10% growth in females raises it by 21.6%. For those over 75, 10% growth in males decreases output by 4.9%, while 10% growth in females raises it by 14.4%. Growth of ten percentage points in the share of days subsidized by Medicaid appears to raise output by only 1.6%, controlling for growth in gender. The 20-year growth in Medicaid share of around 45% thus contributed to 7.2% of the growth in nursing home bed-days, a small share of the over 60% growth in bed-days. It appears that the explosion in Medicaid contributed little to the growth in nursing home bed-days, above and beyond the contribution of gender growth.[18] Medicaid growth appears merely to be responding to differentials in gender growth: since widows are more likely to be poor than married women, unbalanced gender growth raises both demand and the portion of demand subsidized by Medicaid.

If we assume for the moment that 20-year growth in gender and bed-days was distributed evenly across states, we can use the coefficients of the over age 65 regression in the first column of Table 8, along with the data on total growth in Table 7, to decompose the actual 20-year increase in bed-days. Using this method, we compute: the 45.84 percentage point increase in Medicaid increased bed-days by 7.33%; the 59.09% increase in females over 65 increased bed-days by 127.63%; the 48.89% increase

[18] The Medicaid elasticities appear much higher for the within-county regressions reported in Table 8. However, this specification relies exclusively on variation across time, for this 8-year data set, while the within-state specification relies on variation across time and across counties within a particular state. Therefore, we take the within-state estimates to be more reliable. The much higher R^2 of the within-state estimates support this strategy.

in males over 65 decreased bed-days by 61.11%. Observe that the total computed growth, 73.86%, exceeds the actual growth of 64.6%. This should come as no surprise since in fact growth rates will be higher for states with lower initial stocks; in effect, this approximation method will assign excessively high growth rates to larger states and excessively low growth rates to smaller states, so the net bias is likely to be upwards. Even so, however, it is obvious that *the gender effects overwhelmingly dominate the effects of Medicaid expansion.*

We should also suspect in the analysis of Medicaid that the long-run population elasticity will be less than the short-run population elasticity. In Table 9, we find this to be the case.

INSERT TABLE 9 HERE

In the first two columns, we regress the 20 year first-difference in nursing home bed-days on the 20-year first-differences of males and females, and the growth in percentage points of the share of bed-days financed by Medicaid. In the last two columns, we regress the 20-year log-difference in nursing home bed-days on the 20-year log-differences in males and females, and the growth in percentage points of the share of bed-days financed by Medicaid. Once again, we find that the R^2 plummets for the log-difference case; this is probably due, once again, to the loss of sample size (over 15%), and the extremely high growth rates for counties with small elderly and bed-day stocks in 1971. Nonetheless, the coefficient on males continues to be significantly negative and the coefficient on females continues to be significantly positive for both the first-differenced cases and the log-differenced cases. Controlling for Medicaid expansion, we find that the total elasticity of gender-neutral population growth falls to a level between 0.66 and 0.71; roughly speaking, a 10 percentage point increase in the 20-year rate of gender-neutral population growth raises the 20-year rate of bed-day growth by about 7 percentage points. For the over 65 population, about 10 percentage points of this predicted 7 percentage point increase in the rate of bed-day growth can be attributed to a 10 percentage point increase in the growth rate of females, while the remaining decrease of about 3 percentage points can be attributed to the 10 percentage point increase in the growth rate of males. Once again, in the long-run, when people have the opportunity to match again and adjust to transient shocks, elasticities appear lower than in the short-run.

We know that the '70s and '80s did witness an enormous expansion of Medicaid subsidization in the long-term care industry. Of course, by its very nature, this expansion cannot account for the deceleration of nursing home growth during this time period. While our analysis has shown that some of the growth in nursing homes which did take place can be attributed to Medicaid expansion, gender effects appear to dominate quantitatively. Moreover, while Medicaid expansion may have contributed somewhat to acceleration in the growth of nursing home bed-days, changes in growth rates across gender so far represent the only candidate explanation for the sharp *deceleration* in nursing home growth.

4.4 Certificate of Need Laws and the Deceleration of Long-Term Care Growth

This section considers another candidate explanation for the deceleration of nursing home growth. During the late '70s and '80s, a variety of states enacted regulations to curb the growth of the long-term care industry. These regulations took the form of Certificate of Need (CON) laws, which required a justification of nursing home bed "need" as a prerequisite for the construction of new beds. Often the criteria for need rested on the number of beds relative to the number of elderly people. In addition, some states imposed moratoria on bed construction. While there has been a reasonable amount of exit out of and entry into regulation over the past 15 years, the number of states with regulations has been extremely high. In 1978, no states had moratoria, but 40 states had CON laws. By 1986, the number of states with CON laws had climbed to 44, and 10 states had moratoria. By 1994, the number of states with moratoria had climbed to 16, while the number with CON laws had fallen slightly to 41.[19] In this section, we will investigate whether or not CON laws could have contributed to the deceleration of nursing home growth, and whether they explain this deceleration better than the balancing of gender growth.

In particular, we will investigate the following model:

$$LogBed_{it} = \gamma_1 + \gamma_2(LogMale_{it}) + \gamma_3(LogFemale_{it}) + \gamma_4(MedShr_{it}) + \gamma_5(Law) + \varepsilon_{it} \tag{12}$$

Here *Law* is a dummy variable indicating the presence of either a CON law or a moratorium. We continue to control for state fixed-effects. Table 10 presents the results for equation 12, along with other regressions.[20] We find that the presence or absence of legislation has little significant impact on the estimated gender effects.

INSERT TABLE 10 HERE

We immediately see that it will be impossible to identify the exact contribution of CON laws to the slowdown, due to a lack of instruments. The source of bias is obvious: given rational policymaking, states with CON laws will tend to be those with higher levels of beds relative to population. Therefore, the presence of CON laws will be associated with higher bed levels, both across states and even within states, assuming that CON laws become implemented in a given state in response to high bed levels. Moreover, this source of bias cannot be easily offset: since CON laws will be enforced differently across states and even across years, we need not see consistent restraints on output associated with the existence of CON laws. The first four columns of the table present simple regressions which include dummies for CON laws and bed moratoria. The coefficients on the CON law dummies appear positive, most likely as a result of the simultaneity bias discussed. The moratorium dummy appears negatively significant for the over age 75 group. The bias for moratoria policies will be more easily offset, because a moratoria imposes a definite end to bed construction and should thus restrain output similarly across states and across years. However, by comparing the first four columns of Table 10 to the estimates given in Tables 5 and 8, we can see that the estimated gender effects are actually more significant than the estimates we have been considering so far. (Specifically, the male elasticity becomes more negative, while the female elasticity becomes more positive.) Since we have already seen that our previous estimates imply overwhelming gender effects, we can tentatively conclude that legislative differences will not weaken these gender effects.

To test this initial finding further, we examine whether or not the gender elasticities change for states with different legislative practices. In particular, we will split up the sample based on legislative practices and then examine the regression model in equation 10. First, we split up the sample into points at which CON laws obtain, and points at which they do not obtain. We find that CON laws slightly weaken the gender effects, in the sense that they drive the male and female coefficients closer together. However, in both cases, the estimated elasticities lie above those presented in Table 5, while the total elasticity of gender-neutral growth remains unchanged. The magnitude of the weakening is

[19] See Harrington *et al* (1997).

[20] Data on the existence of statewide CON laws and moratoria are presented in Harrington *et al* (1997), for the years 1978, 1982, 1986, 1990, and 1994. We assume that the 1990 data applies to 1991, and use it in conjunction with our 1991 nursing home data.

quite small though. For the over 65 regression, the elasticities fall in absolute value by about .06, while the fall is just .04 for the over 75 regression.

Finally, we look at those observations at which both a CON law and a moratorium apply, and then those observations at which neither apply. It should be noted that while these findings are of interest, they will not exert much influence in the data set as a whole, because each set of observations represents around 5% of the sample. The observations with both a CON law and a moratorium do display a significantly different pattern than the data set as a whole: the total population elasticity rises to around unity, perhaps because in such states a target level of beds per old person is binding; as a result, across such states, one will find that states with higher bed levels must have proportionally higher elderly populations. However, even in the presence of these regulations, we continue to find the qualitative result that growth in men depresses output, and growth in women increases it. In the last two columns of Table 10, we consider gender effects for the unregulated sample. We find that the gender effects are on the higher end of the spectrum: in unregulated markets, gender effects appear to be greater. The imposition of supply restrictions prevents people who would normally be inframarginal consumers from entering nursing homes. As a result, given supply restrictions, increased parity in the marriage market might not lower nursing home utilization, since there exist some people outside of nursing homes who cannot be cared for by a spouse and who will thus enter at the first available opportunity, regardless of the marriage market.

In sum, it appears that the presence of entry barriers may slightly decrease the magnitude of the gender effects, but that even in the presence of legislative barriers to entry, the gender effects do not fall by much. We can conclude that legislative barriers do not vitiate much of the explanatory power possessed by gender effects.

4.5 Changes in Frailty Patterns and Long-Term Care

We will now consider our final candidate for explaining the slowdown in nursing home: changes in the frailty of elderly populations. Unfortunately, frailty data are available for only one year, 1990, from the US Census. As a result, we will not be able to test explicitly the impact of changes in frailty over time. However, since our results hold up for regressions on data from a single year, such changes, while they may explain some part of the change in nursing home growth, will not affect the significance of the gender effects.

In the 1990 Census, respondents were asked if they had a Mobility Limitation (ML) or Self-Care Limitation (SCL). We have data on the number of respondents in the non-institutional population over the ages of 65 and 75, who had one, the other, both, or neither. Therefore, we know the frailty rate in the non-institutionalized population. Unfortunately, this rate is not exogenous and depends heavily on the long-term care market: states with high provision of long-term care will have relatively fewer frail individuals outside of institutions. We need data on the rate of frailty in the entire elderly population, both institutionalized and non-institutionalized.[21] To construct such a measure, we assume that all residents of nursing homes have one or the other limitation, and we assume that they are all over 75. As a result, we construct the total statewide frail population as the statewide number of non-institutionalized individuals with an ML or SCL plus the statewide total number of nursing home

[21] As a result of this endogeneity problem, a wealth of Census data is not usable for us. In particular, the Census breaks down the disability data by marital status and living arrangement. Clearly, however, our model predicts that the rate of marriage in the non-institutionalized population will differ significantly from the rate in the entire population, since states with cheaper long-term care will admit relatively more frail singles and thus will have higher marriage rates in the non-institutionalized population.

residents (from the Area Resource File); the statewide frailty rate is then this total frail population divided by the non-institutional plus institutional populations.

Since our nursing home data from the Area Resource File are reported for 1991, we estimate the following model for 1991:[22]

$$LogBed_i = \gamma_1 + \gamma_2(LogMale_i) + \gamma_3(LogFemale_i) + \gamma_4(Frail_i) + \varepsilon_i$$

Here the variable *Frail* represents the frailty rate, and i corresponds to a county. The results of these regressions are presented in Table 11.

INSERT TABLE 11 HERE

Observe from the first two columns of Table 11 that in the initial regressions, the coefficient on *Frail* is negative. This owes itself to another endogenous component of frailty: poverty. To the extent that lifetime health is endogenous, so is frailty among the elderly, and there will be significant positive correlation between poverty and frailty. Poverty, in addition, will be negatively correlated with nursing home use. To test this hypothesis, we first control for the rate of poverty over the age of 65, and find that the coefficients on frailty drop substantially: for the over 75 age group, the coefficient becomes insignificant, but for the other group, it still remains significantly negative. We then include a dummy for West Virginia, whose Appalachian counties includes several of the nation's poorest, and which is one of the nation's 2 poorest states (next to Arkansas). It is a puzzle that in spite of West Virginia's poverty, its reported rate of poverty among the elderly is not far above the mean. Including a dummy for West Virginia, we find that indeed nursing home use in West Virginia is well below predicted levels and that the coefficients on frailty go to zero. As a final control, we include per capita statewide income. Again, we find that increases in income reduce nursing home use, through gains in health and declines in frailty, and we find that the coefficients on frailty fall further.

Throughout this exercise, however, there appears to be little positive effect of frailty on nursing home use, once we control for population. Even when we introduce controls on poverty, the coefficient on frailty does not become positively significant. Moreover, controlling for frailty, the gender effects remain extremely large, much larger even than the effects we have been using to show that gender imbalance contributed to most of the nursing home slowdown. While variation in frailty across states may in fact contribute to change in nursing home usage, it seems hard to make the case that gender stocks simply proxy for this frailty variation. If anything, it seems that frailty may proxy for changes in gender stocks, since controlling for gender stocks reduces the effect of frailty to insignificance.

[22]Since our nursing home data from the Area Resource File are for 1991, we assume that the 1990 Census data on frailty obtains for 1991 and assume that frailty rates are constant across counties in the same state. Therefore, the county frailty rate is the statewide frailty rate.

5 Concluding Remarks

This paper analyzed theoretically and empirically the way in which the market for long-term care responds to aging. We contrasted our arguments with those claiming that growth in the elderly population induces more than proportional growth in market care. Although economists often argue that aging raises demand for long-term care, we claimed that by raising the availability of substitute home care, aging directly lowers it. Through its covariance with frailty, aging indirectly affects demand, but this may counteract or reinforce the direct negative effect depending on the correlation between longevity and frailty. Furthermore we argued that the importance of marriage in the production of home care implies that gender-specific growth will have a different impact than gender-neutral growth. Indeed, our theoretical and empirical analysis found that growth in males will reduce the output of long-term care. Our theory suggests a natural interpretation of the limited growth in market output which has accompanied the steady growth in the elderly and the rapid growth in the share of US demand subsidized over the last few decades. We were able to account for the rapid deceleration in nursing home growth during the 1980s as a feature of more balanced gender growth during that decade. In our empirical analysis, we found that the majority of the slowdown can in fact be attributed to the sudden balancing of gender growth.

The analysis suggests several avenues of future research. First, we have stressed the importance of the covariation between morbidity and mortality in shaping market responses to aging. Economic theory may offer predictions for this covariance within a cohort at a given point in time and across cohorts over time. In the health human capital model,[23] 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, a negative relationship between frailty durations and mortality will be predicted in this case. Interpreting the duration of frailty in this manner, our analysis suggests that an important question for future research concerns the impact of economic determinants, such as public old-age income or health support, on the covariance between frailty and mortality.[24]

Second, the trade-off between home and market production may be contingent on not only the mean aging of populations but also the *variance* of aging. Eligibility for the receipt of mandatory annuities, Social Security in the US, is age dependent.[25] Mandatory social security subsidizes home care provided by children when the disability of a parent occurs at the retirement ages of those children. 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, in addition to the means, of the parental age distribution matters, because children with relatively older parents receive more subsidies (relative to their income) if they leave the labor market and provide home care for their parents.

[23] See, for instance, Grossman (1972).

[24] Of interest here would be the interrelationship between markets for *short-* and long-term care. Short-term health care may reduce both entry and exit into and out of disability, forces which have offsetting effects on the stock of disabled. The question of how a reduction of death and disability affects the long-term care market seems to raise important general equilibrium issues.

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

These future questions illustrate that which we hope the analysis has demonstrated: the economic aspects of long-term care markets go far beyond the simple financial transactions, on which current economic analysis of long-term care has focused. A deeper and fuller understanding of how these markets respond to aging requires a more substantial economic analysis of how matching and health investments determine the underlying demographics of elderly populations.

6 Appendix

In this appendix, we will prove the propositions stated in the text. The proofs will use the steady-state values from the model of marriage and long-term care laid out in the text:

$$\dot{C} = b - (\varphi_f^{-1} + \varphi_m^{-1})C + \nu_f^{-1}C_f$$

$$\dot{C}_f = \varphi_f^{-1}C - \nu_f^{-1}C_f - \varphi_m^{-1}C_f$$

$$\dot{C}_m = \varphi_m^{-1}C - \nu_m^{-1}C_m - \varphi_f^{-1}C_m$$

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

$$\dot{H}_f = \nu_m^{-1}C_m - \nu_f^{-1}C_f - \varphi_f^{-1}H_f$$

$$\dot{F}_f = \varphi_f^{-1}H_f + (\varphi_m^{-1}C_f + \varphi_f^{-1}C_m) - \nu_f^{-1}F_f$$

$$\dot{F}_m = \varphi_m^{-1}H_m + (\varphi_m^{-1}C_f + \varphi_f^{-1}C_m) - \nu_m^{-1}F_m$$

Proof of Proposition 1

Solving for the steady-state, we find that:

$$H_f = \left(\frac{\varphi_f b}{\varphi_f^{-1} + \varphi_m^{-1} + \frac{\nu_f^{-1}\varphi_f^{-1}}{\nu_f^{-1} + \varphi_m^{-1}}}\right)\left(\frac{\varphi_f}{\varphi_m\varphi_f + \nu_m\varphi_m} - \frac{\varphi_m}{\varphi_f\varphi_m + \nu_f\varphi_f}\right)$$

Therefore, we know that:

$$\varphi_f^{-1}H_f = \left(\frac{b}{\varphi_f^{-1} + \varphi_m^{-1} + \frac{\nu_f^{-1}\varphi_f^{-1}}{\nu_f^{-1} + \varphi_m^{-1}}}\right)\left(\frac{\varphi_f}{\varphi_m\varphi_f + \nu_m\varphi_m} - \frac{\varphi_m}{\varphi_f\varphi_m + \nu_f\varphi_f}\right)$$

It is obvious that this quantity increases in φ_f, but it remains to check that it decreases in φ_m. The term in the second set of parentheses clearly decreases in φ_m, but the first term need not. However, expanding this expression shows that the entire quantity decreases in φ_m. *QED*

Proof of Proposition 2

Solving for the steady-state, we find that the steady-state stocks of frail, unattached males and females satisfy:

$$F_m = \nu_m b\left(\frac{1}{\left(\frac{\varphi_m}{\nu_f} + 1\right)\left(1 + \frac{\varphi_f}{\varphi_m} + \frac{1}{1 + \frac{\nu_f}{\varphi_m}}\right)} + \frac{1}{\left(1 + \frac{\varphi_f}{\nu_m}\right)\left(1 + \frac{\varphi_m}{\varphi_f} + \frac{\varphi_f^{-1}\varphi_m}{1 + \frac{\nu_f}{\varphi_m}}\right)}\right) \tag{13}$$

$$F_f = \frac{\nu_f b}{\left(1 + \frac{\nu_m}{\varphi_f}\right)\left(1 + \frac{\varphi_m}{\varphi_f} + \frac{\varphi_f^{-1}\varphi_m}{1+\frac{\nu_f}{\varphi_m}}\right)} + \frac{\nu_f}{\nu_m} F_m \tag{14}$$

Inspecting equation 13 reveals that an x percent increase in b raises F_m by x percent, while an x percent increase in φ_f, φ_m, ν_f, ν_m also increases F_m by x percent, because it leaves the term in the parentheses unchanged while raising ν_m by x percent. Knowing this about F_m, we can be sure by inspecting equation 14 that an increase in b by x percent will raise F_f by x percent, and that the same will be true for an x percent increase in φ_f, φ_m, ν_f, ν_m. QED

Finally consider the following proposition alluded to in the text:

Proposition 3 *If φ_f and ν_f increase by x percent while all other quantities remain constant, and if F_m is "small," the population of frail single people will increase by more than x percent.*

Proof of Proposition 3

In order to show this formally, we will assume that $F_m = 0$, although clearly the proposition must hold for all F_m small enough relative to F_f. Now suppose that φ_f and ν_f increase by x percent. If $F_m = 0$, then market output is given by $\varphi_f^{-1} H_f$, according to equation 14. Examining that equation, it is clear that $\varphi_f^{-1} H_f$ must rise by more than x percent, because the numerator rises by x percent, and the denominator strictly falls. QED

References

[1] Freedman, V., (1997). "Family Structure and The Risk of Nursing Home Admissions." *Journal of Gerontology: Social Sciences.* 51B.2: S61-S67.

[2] Garber, A., (1994). "To Comfort Always: The Prospects of Expanded Social Responsibility for Long-Term Care." *Individual and Social Responsibility.* V. Fuchs, ed. Chicago and London: The University of Chicago Press.

[3] Gertler, P., (1989). "Subsidies, Quality, and the Regulation of Nursing Homes." *Journal of Public Economics* 39: 33-53.

[4] Gertler, P., and D. Waldman, (1992). "Quality-Adjusted Cost- Functions and Policy Evaluation in the Nursing Home Industry." *Journal of Political Economy* 100: 1232-1256.

[5] *Green Book of Entitlement Programs*, (1995). Ways and Means Committee. House of Representatives. United States Congress.

[6] Grossman, M., (1972). "On the Concept of Health Capital and The Demand for Health." *Journal of Political Economy* 80: 223-50.

[7] Harrington, Charlene, *et al.*, (1997). "State Regulation of the Supply of Long Term Care Providers." *Journal of Applied Gerontology* 16.1 (March 1997): 5-30.

[8] HCFA, (1997). Medicare and Medicaid Statistical Supplement. *Health Care Financing Review.*

[9] HCIA, (1996). *The Guide to The Nursing Home Industry.* Dallas, TX: Arthur Andersen.

[10] Hurd, M., (1990). "Research on The Elderly: Economic Status, Retirement, and Consumption and Saving." *Journal of Economic Literature* 28 (2): 565-637.

[11] Kotlikoff, L., (1989). *What Determines Savings?* Cambridge, MA: MIT Press.

[12] Manton, K., L. Corder, and E. Stallard, (1992). "Estimates of change in chronic disability and institutional incidence and prevalence rates in US elderly population from the 182, 1984, and 1989 Long-term Care Survey." *Journal of Gerontology* 48: S 153-66.

[13] National Center for Social Statistics (NCSS), (1974). Medicaid Recipient Characteristics and Units of Selected Medical Services: 1972. Washington, DC: United States Department of Health, Education, and Welfare.

[14] Norton, E., and J. Newhouse, (1994). "Policy Options for Public Long-Term Care Insurance." *Journal of the American Medical Association* 271(19): 1520-4.

[15] Norton, E., (1995). "Elderly Assets, Medicaid Policy, and Spend-Down in Nursing Homes." *Review of Income and Wealth* 41.3: 309-29.

[16] Norton, E., (1997). "Long-Term Care." *Handbook of Health Economics*, forthcoming. J. Newhouse and A. Culyer, Eds. New York: North-Holland.

[17] Office of National Health Statistics. (1994). Office of the Actuary: National Health Expenditures, 1993. *Health Care Financing Review* 16.1. HCFA Publication No. 03361. Health Care Financing Administration. Washington, DC: US Government Printing Office.

[18] Office of National Health Statistics. Office of the Actuary: National Health Expenditures, 1987. *Health Care Financing Review* 10.2. HCFA Publication No. 03276. Health Care Financing Administration. Washington, DC: US Government Printing Office.

[19] Philipson, T., (1998). "Perfect Substitution of Organizational Form; The Case of Long-Term Care." *Non-Profit Hiospitals*, forthcoming. D. Cutler, Ed. National Bureau of Economic Research. Chicago: University of Chicago Press.

[20] Philipson, T., and G., Becker (1998). "Old-Age Longevity and Mortality Contingent Claims." *Journal of Political Economy* 106 (3): 550-74.

[21] Posner, R. A., (1995). *Aging and Old Age.* Chicago: University of Chicago Press.

[22] Ruther, M., Reilly, T., Silverman, H., and D. Aibott, (1991). *Medicare and Medicaid Data Book.* Baltimore, MD: Health Care Financing Administration.

[23] Schoeni, R., and K. McGarry, (1995). "Transfer Behavior in the Health and Retirement Study: Measurement and the Redistribution of Resources Within the Family." *Journal of Human Resources* 30: S184-S226.

[24] Stolnitz, G., (1956). "A Century of International Mortality Trends." *Population Studies* 10: 17-42.

[25] United Nations, (1990). *Patterns of First Marriage: Timing and Prevalence.* New York: United Nations.

[26] United States Department of Health and Human Services. National Center for Health Statistics. *National Nursing Home Survey.* 1973-1974, 1977, 1985, 1995 [Computer File]. Hyattsville, MD: U.S. Dept. of Health and Human Services. National Center for Health Statistics [producer]. 1988. Ann Arbor, MI: Inter-university Consortium for Political and Social Research [distributor]. 1992.

[27] Weisbrod, B., (1988). *The Non-Profit Economy.* Cambridge, MA: Harvard University Press.

[28] Wise, D., (1992). *Topics in The Economics of Aging.* National Bureau of Economic Research. Chicago: University of Chicago Press.

Table 1: Summary Statistics for Gender and Statewide U.S. Long-Term Care Output (1989-94).[a]

	Levels		5-Year
	Mean	Std Deviation	Growth[b]
Yearly Bed Days (1000)[c]	11148.18	10337.91	21.94%
Male Population Over 65 (1000)	253.35	276.18	10.18%
Female Population Over 65 (1000)	373.74	401.53	8.33%
Male Population Over 75 (1000)	95.09	104.43	16.20%
Female Population Over 75 (1000)	172.57	182.90	12.95%
Male Population Over 85 (1000)	17.77	19.49	20.48%
Female Population Over 85 (1000)	45.92	48.13	20.84%
Total Population Over 65 (1000)	627.09	677.26	9.06%
Share of Females Over 65	0.59	0.02	-0.41% †
Total Population Over 75 (1000)	267.66	285.61	14.06%
Share of Females Over 75	0.64	0.03	-0.64% †
Total Population Over 85 (1000)	63.69	67.15	20.73%
Share of Females Over 85	0.72	0.03	0.06% †

Source: HCIA (1996), *Guide to the Nursing Home Industry*.

[a]All Data are collected at the state level.

[b]Refers to growth rates averaged across states.

[c]Calculated as (Total Beds)*365*(Occupancy Rate), where Occupancy Rate is the median statewide occupancy rate. Data available from 1991-3; 5-year Growth rate extrapolated from 2-year Growth Rate assuming constant annual growth rates.

†Refers to the absolute change in percentage points.

Table 2: Summary Statistics for Gender and U.S. National Long-Term Care Output (1989-94).[a]

	Levels		5-Year
	1989	1994	Growth[b]
Yearly Bed Days (1000)[c]	554917.42	595616.64	7.33%
Male Population Over 65 (1000)	12333.00	13491.96	9.40%
Female Population Over 65 (1000)	18348.63	19670.30	7.20%
Male Population Over 75 (1000)	4509.22	5204.46	15.42%
Female Population Over 75 (1000)	8308.35	9261.74	11.47%
Male Population Over 85 (1000)	831.13	992.57	19.42%
Female Population Over 85 (1000)	2136.60	2548.70	19.29%
Total Population Over 65 (1000)	30681.63	33162.26	8.09%
Share of Females Over 65	0.60	0.59	-0.49% [†]
Total Population Over 75 (1000)	12817.57	14466.19	12.86%
Share of Females Over 75	0.65	0.64	-0.80% [†]
Total Population Over 85 (1000)	2967.73	3541.27	19.33%
Share of Females Over 85	0.72	0.72	-0.02% [†]

Sources: HCIA (1996), *Guide to the Nursing Home Industry*. *National Nursing Home Survey* (1985).

[a]Except where indicated, all Data are aggregated from the state level.

[b]Refers to nationwide growth rates.

[c]HCIA Data available from 1991-3. Using 1993 HCIA data and 1995 NNHS data, 1989 and 1994 data are extrapolated under assumption of constant annual growth rates between 1985 and 1994.

[†]Refers to the absolute change in percentage points.

Table 3: Effect of Gender-Specific Stocks of Old on U.S. Statewide Long-Term Care Output (1991-94).[a]

	Log Yearly Bed Days[b]					
	Levels			First Differences[c]		
Constant	3.73 [†] 16.33	4.26 [ǀ] 27.02	5.42 [†] 35.31	0.01 1.56	0.01 1.19	0.00 0.16
Log Males Over 65	-1.93 [†] -5.19			-0.56 -0.52		
Log Females Over 65	2.85 [†] 8.08			1.55 1.51		
Log Males Over 75		-1.17 [†] -6.62			-0.22 -0.41	
Log Females Over 75		2.10 [ǀ] 12.77			1.01 [ǀ] 1.70	
Log Males Over 85			-0.57 [ǀ] -5.31			0.17 0.57
Log Females Over 85			1.53 [†] 14.91			0.42 1.14
R-Squared	0.93	0.95	0.96	0.04	0.03	0.02
Number of Observations	203	203	203	152	152	152

[a]Robust T-statistics given below point estimates.

[b]Yearly Bed Days defined as (Total Beds)*365*(Occupancy Rate), where Occupancy Rate is the median statewide occupancy rate.

[c]For logarithmic variables, the difference in logarithms is used.

[ǀ]Significantly different from zero with 95% confidence.

[†]Significantly different from zero with 90% confidence.

Table 4: Summary Statistics for Gender and Countywide U.S. Long-Term Care Output (1971-91).[a]

	1991[b]		1971[b]		20-Year
	Mean	Std Deviation	Mean	Std Deviation	Growth[c]
Yearly Bed Days (1000)[d]	234.22	660.26	142.30	500.38	138.87%
Nursing Home Residents (1000)	0.58	1.60	0.35	1.19	146.44%
Yearly Medicaid Bed Days (1000)	7595.03	7187.35	2584.04	3739.44	1728.23%
Share of Medicaid Bed Days	0.70	0.11	0.27	0.19	42.99% ‡
Female Population Over 65 (1000)	6.17	18.71	3.90	13.61	66.26%
Male Population Over 65 (1000)	4.15	12.33	2.80	9.16	47.34%
Share of Females Over 65	0.59	0.03	0.56	0.04	3.13% †
Total Population Over 65 (1000)	10.32	31.02	6.70	22.76	57.38%
Female Population Over 75 (1000)	2.83	8.50	1.57	5.41	93.65%
Male Population Over 75 (1000)	1.54	4.56	0.99	3.17	57.16%
Share of Females Over 75	0.63	0.04	0.58	0.06	5.22% †
Total Population Over 75 (1000)	4.36	13.03	2.56	8.57	76.90%

Sources: United States Bureau of Health Professions Area Resource File (1940-1995). 1971 Medicaid Data from NCSS (1974). 1991 Medicaid Data from HCIA (1996).

[a]Data collected at countywide level, except for Medicaid Data, which is collected by state.

[b]Population data linearly interpolated using census data from 1970, 1980, and 1990. Medicaid Data reported for 1971 comes from 1972.

[c]Refers to Growth Rates Averaged Across Counties. Excludes counties with zero 1971 values: 2 for census variables, and 441 for long-term care variables. Growth rates for population variables calculated using raw 1970 and 1990 data.

[d]Calculated as (Total Beds)*365.

†Refers to the absolute change in percentage points.

‡Refers to the absolute change in percentage points averaged across states.

Table 5: Short-Run Effects of Gender-Specific Stocks of Old on U.S. Countywide Long-Term Care Output (1971-91).[a]

	Log Yearly Bed Days[b]							
	Within States				Within States, Years		Within Counties	
Constant	3.85 †	4.51 †	3.83 †	4.52 †	3.92 †	4.49 †	2.62 †	4.27 †
	134.85	185.61	132.78	24.13	130.22	166.72	18.87	42.42
Log Males Over 65	-1.33 †		-1.23 †		-1.27 †		-0.87 †	
	-31.98		-27.96		-28.70		-13.45	
Log Females Over 65	2.25 †		2.15 †		2.18 †		1.97 †	
	57.17		51.45		52.21		36.01	
Log Males Over 75		-0.48 †		-0.53 †		-0.56		-0.06
		-17.97		-18.11		-18.97		-1.56
Log Females Over 75		1.44 †		1.48 †		1.51 †		1.08 †
		57.71		54.57		55.04		37.33
Time[c]			0.004 †	-0.002 †				
			6.32	-3.88				
R-Squared	0.78	0.79	0.78	0.79	0.78	0.79	0.27	0.29
Number of Observations	21948	21948	21948	21948	21948	21948	21948	21948

[a]Robust T-statistics given below point estimates.

[b]Yearly Bed Days defined as (Total Beds)*365.

[c]Refers to number of years elapsed since 1971.

[†]Significantly different from zero with 95% confidence.

Table 6: Long-Run Effects of Gender-Specific Stocks of Old on U.S. Countywide Long-Term Care Output (1971-91).[a]

	Yearly Bed Days[b]			
	First Differences[c]		Log Differences[c]	
Constant	16029.37 †	16890.11 †	0.21 †	0.10 †
	4.59	5.36	7.50	3.18
Males Over 65	-75.50 †		-0.40 †	
	-5.63		-2.60	
Females Over 65	78.38 †		1.20 †	
	8.26		7.24	
Males Over 75		-71.98 †		-0.10
		-3.82		-1.05
Females Over 75		90.45 †		0.94 †
		8.56		9.62
R-Squared	0.77	0.78	0.11	0.13
Number of Observations	3082	3082	2610	2610

[a]Robust T-statistics given below point estimates.

[b]Yearly Bed Days defined as (Total Beds)*365.

[c]Refers to both dependent and independent variables.

†Significantly different from zero with 95% confidence.

Table 7: Summary Statistics for Gender and U.S. National Long-Term Care Output (1971-91).[a]

	Levels[b]		20-Year
	1971	1991	Growth
Yearly Bed Days (100,000)[c]	4385.83	7218.60	64.59%
Nursing Home Residents (100,000)	10.76	17.94	66.76%
Medicaid Bed Days (100,000)	1059.46	4389.35	314.30%
Share of Medicaid Bed Days	0.24156	0.70	45.84% †
Female Population Over 65 (100,000)	119.807	190.60	59.09%
Male Population Over 65 (100,000)	86.08	128.1684	48.89%
Share of Females Over 65	0.58	0.60	1.60% †
Total Population Over 65 (100,000)	205.89	318.77	54.83%
Female Population Over 75 (100,000)	48.09	87.55	82.04%
Male Population Over 75 (100,00)	30.55	47.48	55.40%
Share of Females Over 75	0.61	0.65	3.69% †
Total Population Over 75 (100,00)	78.65	135.03	71.69%

Sources: Nursing Home Data from Area Resource File (1940-1995). Population Data from United States Census (1970, 1980, and 1990). 1971 Medicaid Data from NCSS (1974). 1991 Medicaid Data from HCIA (1996).

[a]Medicaid data aggregated from statewide statistics. All other data aggregated from countywide statistics.

[b]Assuming constant 10 year growth rates, population data interpolated using Census data from 1970, 1980, and 1990.

[c]Calculated as (Total Beds)*365.

†Refers to the absolute change in percentage points.

Table 8: Short-Run Impacts of Medicaid on Gender Determinants of U.S. Countywide Long-Term Care Output (1971-91).[a]

	Log Yearly Bed Days[b]					
	Within States				Within Counties	
Constant	3.88 †	4.60 †	3.88 †	4.55 †	5.05 †	5.72 †
	127.35	178.42	123.01	166.76	28.14	39.22
Log Males Over 65	-1.25 †		-1.25 †		-0.32 †	
	-28.52		-28.06		-4.31	
Log Females Over 65	2.16 †		2.16 †		1.09 †	
	51.95		51.02		14.66	
Log Males Over 75		-0.49 †		-0.52 †		0.01
		-16.97		-17.64		0.37
Log Females Over 75		1.44 †		1.47 †		0.78 †
		53.33		53.18		19.65
Statewide Share of Medicaid Bed-Days	0.16 †	-0.03	0.17 †	0.16	0.55 †	0.37 †
	7.01	-1.17	3.78	3.65	18.18	11.60
Time[c]			0.000	-0.006 †		
			-0.24	-4.98		
R-Squared	0.79	0.79	0.79	0.79	0.29	0.29
Number of Observations	20011	20011	20011	20011	20011	20011

[a]Robust T-statistics given below point estimates.
[b]Yearly Bed Days defined as (Total Beds)*365.
[c]Refers to number of years elapsed since 1971.
†Significantly different from zero with 95% confidence.
‡Significantly different from zero with 90% confidence.

Table 9: Long-Run Impacts of Medicaid Subsidies on Long-term care growth in U.S. Counties (1971-91).[a]

	Yearly Bed Days[b]			
	First Differences[c]		Log Differences[c]	
Constant	10930.04	19149.04 †	0.14 †	0.05
	1.37	5.49	3.02	1.08
Males Over 65	-75.75 †		-0.32 †	
	-5.53		-1.98	
Females Over 65	78.44 †		0.98 †	
	8.03		5.56	
Males Over 75		-77.58 †		-0.07
		-3.93		-0.70
Females Over 75		92.62 †		0.78 †
		8.31		7.47
Growth in Medicaid	14956.69	-5.89	0.28 †	0.25 †
Share[d]	1.08	-0.40	3.90	3.53
R-Squared	0.76	0.78	0.09	0.10
Number of Observations	2431	2431	2065	2065

[a]Robust T-statistics given below point estimates. All differences are taken over the entire 20 year period.

[b]Yearly Bed Days are (Total Beds)*365.

[c]Except where noted, refers to dependent and independent variables.

[d]Simple 20-year difference of shares.

†Significantly different from zero with 95% confidence.

Table 10: Effects of State Legislation on U.S. Countywide Long-Term Care Output (1971-91).[a]

	Log Yearly Bed Days[b]											
	Full Sample				States with CON		States Without CON		States with CON and Moratorium		States with Neither CON nor Moratorium	
Constant	3.60 † 80.59	4.36 † 110.46	3.69 † 87.25	4.45 † 120.01	3.81 † 94.91	4.43 † 127.50	3.94 † 46.30	4.54 † 60.17	3.26 † 25.87	3.96 † 35.81	3.68 † 27.09	4.34 † 37.15
Log Males Over 65	-1.39 † -23.04		-1.39 † -23.04		-1.49 † -23.19		-1.55 † -10.62		-0.99 † -5.89		-1.42 † -6.19	
Log Females Over 65	2.29 † 39.89		2.30 † 39.82		2.40 † 39.28		2.45 † 17.46		1.98 † 12.48		2.34 † 10.74	
Log Males Over 75		-0.73 † -16.78		-0.74 † -17.02		-0.78 † -16.94		-0.74 † -6.90		-0.40 † -3.11		-0.83 † -5.03
Log Females Over 75		1.66 † 40.72		1.68 † 40.88		1.73 † 39.86		1.67 † 16.52		1.43 † 11.99		1.77 † 11.51
Statewide Share of Medicaid Bed-Days	0.43 † 11.28	0.17 † 4.66	0.42 † 10.20	0.18 † 4.46								
CON Law Present[c]	0.10 † 6.50	0.09 † 6.33										
Bed Moratorium[c] Present			-0.02 -1.14	-0.04 † -2.23								
R-Squared	0.81	0.82	0.81	0.82	0.80	0.82	0.81	0.82	0.75	0.76	0.76	0.77
Number of States	51	51	51	51	51	51	19	19	11	11	15	15
Number of Observations	10273	10273	10273	10273	9413	9413	1782	1782	1491	1491	1040	1040

Source: Certificate of Need and Moratorium data for 1978, 1982, 1986, and 1990, from Harrington et al (1997).

[a]T-statistics given below point estimates based on White's heteroskedasticity-consistent variance estimator. All regressions are within-state.

[b]Yearly Bed Days defined as (Total Beds)*365.

[c]Dummy variable which takes on the value one when the specified condition is met.

†Significantly different from zero with 95% confidence.

Table 11: Effects of Cross-State Variation in Frailty on U.S. Countywide Long-Term Care Output (1991).[a]

	Log Yearly Bed Days[b]							
Constant	5.19 † 51.45	5.51 † 48.43	4.90 † 41.79	5.09 † 39.28	4.77 † 40.99	5.01 † 38.64	9.40 † 8.09	9.19 † 8.32
Log Males Over 65	-1.61 † -14.20		-1.56 † -13.78		-1.51 † -13.45		-1.52 † -13.65	
Log Females Over 65	2.50 † 22.38		2.43 † 21.75		2.38 † 21.47		2.40 † 21.75	
Share of Frail Persons Over Age 65[c]	-4.66 † -11.92		-1.83 † -2.88		-0.927 -1.488		-0.11 -0.18	
Log Males Over 75		-0.70 † -8.48		-0.68 † -8.24		-0.66 † -8.04		-0.67 † -8.21
Log Females Over 75		1.64 † 20.19		1.60 † 19.85		1.58 † 19.68		1.60 † 19.88
Share of Frail Persons Over Age 75[c]		-2.36 † -8.16		-0.32 -0.79		0.00 -0.01		0.54 1.33
Rate of Poverty for Persons Over Age 65			-1.97 † -6.69	-1.93 † -8.21	-2.315 † -8.019	-2.04 † -8.72	-3.48 † -8.89	-3.05 † -9.25
West Virginia[d]					-0.486 † -4.823	-0.47 † -4.78	-0.58 † -5.63	-0.55 † -5.46
Log Statewide Income Per Capita							-0.48 † -4.03	-0.44 † -3.86
R-Squared	0.85	0.87	0.86	0.87	0.86	0.87	0.86	0.87
Number of Observations	2984	2894	2894	2894	2894	2894	2894	2894

Sources: Frailty data compiled from Area Resource File and U.S. Census (1990).
Income data from HCIA (1996).
[a]Robust T-statistics given below point estimates.
[b]Yearly Bed Days defined as (Total Beds)*365.
[c]Calculated as the statewide non-institutionalized population over the given age with a Self-Care Limitation or Mobility Limitation (1990 U.S. Census), plus the statewide nursing home population (Area Resource File), divided by the statewide non-institutionalized population over the given age (1990 U.S. Census), plus the statewide nursing home population.
[d]Dummy variable for counties in West Virginia.
†Significantly different from zero with 95% confidence.