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THE BEST OF TIMES, THE WORST OF TIMES:
UNDERSTANDING PRO-CYCLICAL MORTALITY

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ABSTRACT

A growing literature documents cyclical movements in mortality and health. We examine this pattern more closely and attempt to identify the mechanisms behind it. Specifically, we distinguish between mechanisms that rely on fluctuations in own employment or time use and those involving factors that are external to the individual. Our investigation suggests that changes in individuals' own behavior contribute very little to pro-cyclical mortality. Looking across broad age and gender groups, we find that own-group employment rates are not systematically related to own-group mortality. In addition, we find that most of the additional deaths that occur during times of economic growth are among the elderly, particularly elderly women, who have limited labor force attachment. Focusing on mortality among the elderly, we show that cyclicality is especially strong for deaths occurring in nursing homes, and is stronger in states where a higher fraction of the elderly reside in nursing homes. We also demonstrate that staffing in skilled nursing facilities moves counter-cyclically. Taken together, these findings suggest that cyclical fluctuations in the mortality rate may be largely driven by fluctuations in the quality of health care.

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I. Introduction

Why do death rates rise when the unemployment rate falls? Pro-cyclical mortality rates in the United States (and elsewhere) are now well-documented, but the causes of this association remain poorly understood. In this paper, we show that pro-cyclical mortality in the United States is not driven by changes in individuals' own time use associated with their own cyclical employment changes. Instead, we point to alternative mechanisms, including cyclical changes in the quality of health care as a major driver of overall pro-cyclical mortality.

The finding of pro-cyclical mortality has been established in a series of influential papers by Christopher Ruhm (2000, 2003, 2005a,2005b), based on estimates that exploit state-year panel data, and include state and year fixed effect and state-specific time trends. A typical estimate (Ruhm 2000) suggests that a one-percentage point increase in a state's unemployment rate leads to a 0.54% reduction in that state's mortality rate. When applied to U.S. mortality counts from 2006, this implies that a one percentage point increase in unemployment would lead to about 13,000 fewer annual deaths.

Ruhm's findings are widely cited in the health economics literature and have been echoed in work by Dehejia and Lleras-Muney (2004) who find that infant health outcomes and economic downturns are positively linked. The most common interpretation has been that good economic times have a negative impact on individuals' health because of an increase in the opportunity cost of time, and the resulting changes in individuals' decisions about how to allocate their time. Ruhm (2005b), for example, finds that obesity and smoking both exhibit a pro-cyclical pattern, and that diet and exercise also improve when the unemployment rate rises – patterns that are consistent with changes in the value of time associated with work.[1] However, pro-cyclical mortality could also be driven by other mechanisms that have not been fully explored.

The purpose of this paper is to shed light on those mechanisms. We are particularly interested in separating the effects of changes in individual behavior that result from changes in one's own employment status from the effects of other factors that fluctuate with the unemployment rate. This distinction is important because of differences in the associated policy prescriptions, but understanding the underlying mechanisms will also shed light on a well-known empirical puzzle--while mortality rates are pro-cyclical, job loss is known to have negative effects on individuals' health. Sullivan and von Wachter (2009), for example, find that individuals who experience a job loss via a mass-layoff experience a substantial increase in their mortality hazard that lasts over the next 20 years. As Ruhm (2008) notes, the estimated impact of individuals' own job loss can be reconciled with the aggregate patterns only if the aggregate fluctuations in mortality are not concentrated among those who change employment status. This suggests that the mechanisms driving pro-cyclical mortality are more complex than a simple connection between own-employment and health.

In the next section, we describe our data and econometric methodology. In section III, we present our results, starting first with a replication and extension of Ruhm's basic analyses (Ruhm, 2000), and then disaggregating by age, gender, and cause of death. Our regressions produce results that are consistent with earlier findings but suggest a previously neglected set of underlying mechanisms. We further explore these possibilities in the remainder of section III, and conclude in section IV.

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[1] Reference to changes in health behaviors as the primary (or only) mechanism behind pro-cyclical mortality is especially common in the news media. See for example, New York Times articles, "Good Economics Times Can Mean Health Risk." May 30, 2005, or "Are Bad Times Healthy?" October 6, 2008.

II. Data and Methodology

We begin by replicating Ruhm's analysis with his own data which he generously shared with us. The basic regression equation takes the following form:

$$H_{jt} = \alpha_t + X_{jt}\beta + E_{jt}\gamma + S_j + S_jT + \varepsilon_{jt} \quad (1)$$

where H is the natural log of the mortality rate in state j and year t, E is a measure of the state's economic health (usually the state unemployment rate), X is a vector of demographic controls including the fraction of the population who are: less than five years old, greater than 65 years old, high school dropouts, with some college, college graduates, black and Hispanic. Most of Ruhm's control variables come from the Census decadal counts and are interpolated in between Census years. The vector of year specific fixed effects, α_t, captures national time effects, and the vector of state specific indicator variables, S_j, controls for time-invariant state characteristics. State-specific time trends are also included. State unemployment rates are taken from unpublished statistics put together by the Bureau of Labor Statistics, and mortality rates come from Vital Statistics publications. The analysis is based on data from 1972-1991.

Our replication results are presented in the first column of Table 1. We present estimates produced by both unweighted regressions and regressions weighted by state-year population,[2] although we find that weighting makes little difference in the magnitude of the estimated coefficient on the state unemployment rate, which is between -0.0054 and -0.0056. The estimates, which are nearly identical to Ruhm's, suggest that a one percentage point increase in the unemployment rate is associated with a 0.5 percent decrease in the predicted death rate.

[2] The two choices of weights are motivated by distinct conceptual questions. Using population weights is appropriate to estimate the degree to which economic conditions contribute to overall fluctuations in U.S. mortality. On the other hand, the un-weighted regressions address the impact on a typical state's mortality rate.

In order to investigate the potential mechanisms behind pro-cyclical mortality, we will exploit both new data sources and additional years of data. The remaining columns of Table 1 show what happens to the estimated relationship between mortality and unemployment as we systematically make these changes. Column 2 shows what happens when we continue to use Ruhm's data but eliminate years between 1972 and 1977. Ultimately, we want to extend our analysis through 2006 so that we can include more recent business cycles in our analysis, but we do not have a consistent measure of the unemployment rate between 1972 and years beyond 2000. Instead, we pool monthly CPS files to construct employment and unemployment rates by state and demographic group between 1978 and 2006 (estimates for all states prior to 1978 are not available in the CPS). Here, we show that eliminating the first six years of Ruhm's data has little impact on the estimated coefficient; the estimated effect of a one percentage point rise in the unemployment rate continues to be approximately -0.005.

Column 3 shows how the estimates change when we replace Ruhm's mortality rate variable with a "new and improved" measure of the mortality rate whose numerator is based on death counts from Vital Statistics' micro-record "multiple cause of death" files and whose denominator comes from population counts collected by the National Cancer Institute's Surveillance Epidemiology and End Results (Cancer-SEER) program. The "multiple cause of death" data are less aggregated than the Vital Statistics data Ruhm uses, and this will later allow us to construct state-level death counts by narrowly defined age groups. We replace Ruhm's population estimates with the Cancer-SEER population counts because the Cancer-SEER estimates are based on an algorithm that incorporates information from Vital statistics, IRS migration files and the Social Security database. As such, they are likely to be more accurate than population estimates that are interpolated between Census years. Changing the dependent variable reduces the estimated unemployment effect by about 20% (from -0.005 to -0.004), but it continues to be strongly statistically significant. Most of this change is driven by the change in the population denominator.

In the fourth column of Table 1, we replace Ruhm's unemployment variable with the CPS unemployment rate. We also replace some of Ruhm's control variables, which are interpolated between census years, with state-year measures of the same variables calculated from the CPS data. We add in a richer set of covariates to control of the state's age distribution. These changes have little effect on the estimated unemployment effect when the regressions are weighted, although they do increase the magnitude of the estimate in the unweighted regressions from -.004 to -.005.

Next, we extend the data through 2006 (column 5). We find that adding fifteen years of data cuts the estimated coefficient on the unemployment rate in half: the new coefficient estimate is between -0.002 (weighted) and -0.003 (unweighted), which suggests that the overall effect of the business cycle on mortality may not be as large as previously thought. On the other hand, the smaller coefficient estimates may also result from other important changes that have occurred over the past fifteen years. In particular, there have been remarkable increases in longevity —between 1978 and 2006, for example, the fraction of Americans over age 65 grew from 11 to 12.5 percent. If this shift in the age distribution occurred unevenly across states, then, given the tight correlation between mortality and age, controlling for these shifts could prove to be very important. Indeed, the age structure in different parts of the country has evolved quite differently over this time period. In California, the fraction of individuals over age 65 increased by less than a percentage point, from 10.0 to 10.8 percent, but in Michigan the fraction of residents over age 65 increased by 3 percentage points, from 9.5 to 12.5 percent.

We control for this phenomenon by replacing the dependent variable with the log of an age-adjusted mortality rate. Consider the mortality rate for state j in a given year t, and note that it can be written as the sum of each age-specific mortality rate weighted by the fraction of individuals in each age interval

$$MR_{jt} = \sum_{a=0}^{85+} MR_{ajt} f_{ajt}$$

In order to abstract from within state-year changes in f_{ajt}, we replace the variable with the 1990 nationwide fraction of individuals in each age category, $f_{a\text{-}US\text{-}1990}$. This creates a measure of the state-year mortality rate that holds the age distribution constant and is defined only by the state-year cell's relative number of deaths. Figure 1, which plots our age-adjusted and unadjusted mortality rates over time, suggests the potential importance of this adjustment; because the U.S. population is aging, the unadjusted series appears to be relatively flat, while the age-adjusted series shows a fairly dramatic decline over time.

Replacing the unadjusted mortality variable with an age adjusted mortality rate turns out to have important effects on our estimates. In column 6, the estimated coefficient on the unemployment rate moves back up to -0.0033. State-specific shifts in the age distribution are clearly correlated with state-level unemployment movements over this period. Note that the need to age-adjust our dependent variable is directly related to our inclusion of additional years of data, which creates a longer period over which states' age distributions can evolve differentially. If we age adjust the mortality rate and repeat the analysis only for the years 1978 through 1991, the estimated unemployment coefficient only moves from -0.0038 to -0.0043 (not shown). [3]

Taken as a whole, our changes have a limited impact on the estimated association between macroeconomic fluctuations and health. Consistent with Ruhm's earlier studies, every entry in Table 1 is negative, statistically significant, and of substantive magnitude. In the remainder of the paper we will focus on weighted regressions in which the dependent variable is the age adjusted mortality rate.

[3] In earlier work (Miller, et al., 2009) we estimate the coefficient on the unemployment rate to be approximately -.005. This estimate is based on data ending in 2004, and we verify that it is sensitive to which years are included. In general, we find that the estimated coefficient on the unemployment rate declines as we add additional years of data after 2000.

III. Why are Recessions Good for your Health?

The previous analyses confirm that mortality rates exhibit a pro-cyclical pattern, and that this pattern persists through the early 2000s. The question that we wish to investigate is *why* the probability of dying increases when economic times are good. Ruhm (2000) proposes four possible mechanisms. First, leisure time declines when the economy improves, making it more costly to undertake health-producing activities that are time-intensive. Second, health may be an input into the production of goods and services. Hazardous working conditions, job related stress and the physical exertion of employment, for example, may all have negative effects on health, and are expected to increase when the economy is expanding. Both of these mechanisms work through changes in individuals' own work hours or opportunity cost of time. A third set of explanations is that the relationship reflects the impact of external factors that fluctuate with the economy. For example, when more people are working, roadways are more congested, and this leads to an increase in the probability of being involved in a fatal auto accident. Evans and Graham (1988) and Ruhm (1995) show that drinking and driving exhibit a pro-cyclical pattern, and Ruhm (2000) shows that motor vehicle fatalities are more sensitive to the business cycle than any other cause of death. Similarly, pollution may vary over the business cycle and contribute to mortality fluctuations.[4] Ruhm's fourth hypothesis is that business cycles affect geographic mobility, which may increase crowding or otherwise bring transition costs that impact mortality. A final possibility raised here is that labor market changes leading to changes in the quantity or quality of purchased inputs to health (such as health care workers), could also affect mortality rates. This mechanism would also be external to

[4] Chay and Greenstone (1999) have established an important connection between pollution and cyclical mortality. This effect seems to operate primarily through infant mortality, however. Chay, Dobkin and Greenstone (2003) show little evidence that pollution has powerful effects on adult mortality. We argue below that only mechanisms having a substantial effect on mortality at older ages are likely to play a substantive role in explaining the overall relationship between mortality and the business cycle. Nevertheless, the role of pollution (including cyclical changes in pollution) on mortality at older ages deserves further study.

individual decisions and outcomes. The estimates that we present in the rest of our paper lead us to conclude that this mechanism may be critical.

III.A. Mortality patterns by age and gender

We begin our investigation of these potential mechanisms by re-estimating equation (1) separately for subgroups defined by gender and age. Estimates for men and women are presented in the lower panels of Table 1. Focusing on our preferred specification, in column 6, we find that while a one percentage point increase in the unemployment rate decreases the female mortality rate by 0.4 percent, it decreases the male mortality rate by only 0.25 percent. This gender difference only emerges when we add the later years of data, however. In column (4), where we focus on the years of data analyzed by Ruhm (2000), the estimated impact on men is actually larger than it is for women. This suggests that the factors that are driving the pro-cyclical pattern may be changing over time. A similar pattern emerges when we repeat our analyses among those 65 and older. Estimated coefficients based on data through 2006 are larger for women than for men. Because older women have particularly low labor force attachment, this result is our first indication that mechanisms independent of individuals' own employment and time use may be at work.

Next, we use our detailed mortality data to estimate equation (1) separately for age groups defined by one- or five-year intervals. Figure 2 displays the unemployment rate coefficients for each year of age. The largest coefficient estimates are among the very young, while coefficients for working-age groups are close to zero, and often positive. Beyond age 60, the estimated coefficients again turn negative.

To reduce the number of estimates while still highlighting the main patterns, Table 2 shows comparable estimates for five-year age groups. Like Ruhm, we find that the mortality rates of young adults are more sensitive to the business cycle than are the mortality rates of most other groups. For example, we estimate that a one percentage point increase in the unemployment rate reduces the mortality rate among 20-24 year olds by about 2 percent. Ruhm (2000) estimates that a one percentage point increase in the unemployment rate reduces the mortality rate among 20-44 year olds by a similar amount. Since young adults' employment fluctuates more than other workers', this finding at first appears to be consistent with the hypothesis that pro-cyclical declines in health are driven by changes in individuals' own behavior. Closer inspection of Table 2, however, reveals several patterns that are not consistent with this story. First, while we estimate a large semi-elasticity among 20-24 year olds, estimates for individuals between ages the ages of 25-59 are substantially smaller. Indeed, for many prime-age workers, the point estimate is near zero. Second, some of the biggest coefficient estimates in our table are associated with age groups that are certain not to be working, such as 0-4 year olds. Finally, the coefficient estimates among those who are over 65 tend to be more negative than the estimates among those 35-64.

Previous work by Ruhm (2007) makes clear that examination of the subgroup coefficient estimates alone is not sufficient to understand what is driving overall cyclicality. This is because deaths are very unevenly distributed across age groups, and only those age groups where a substantial number of deaths occur will make a large contribution to the overall unemployment rate coefficient. For example, even though the coefficient estimates are largest among young people this will not have much impact on fatalities overall because deaths among children and adolescents are rare. To understand how different age groups contribute to the overall estimates in Table 1 we need to weight the age-specific coefficients by the number of deaths in each age group. Columns 4 through 6 show the predicted number of additional deaths that will result from a 1 percentage point decrease in the unemployment rate, using the number of deaths that occurred within each age group in 2006.[5] We see immediately that most of these additional deaths occur among those with relatively weak labor force attachment: fewer than 10% of the additional deaths occur among those between the ages of 25 and 64.[6] In fact, we predict that improvements in the unemployment rate lead to more additional deaths among 0-4 year olds than among 30-50 year olds. In contrast, 70% of the additional deaths are among those over age 70. These results strongly suggest that the mechanisms at play must go beyond changes in individuals' own work behavior.

The remainder of Table 2 breaks down our estimates by gender. The overall point estimate for women (-0.0035) is larger than the estimated coefficient for men (-0.0022), echoing our findings in Table 1. Cyclical mortality appears to be particularly strong among elderly women. The coefficient estimates are notably larger for elderly women than they are for men in the same age range, and women age 65+ account for 55% of the roughly 6700 additional deaths (across all ages and genders) that are predicted to result from a 1 percentage point drop in unemployment. In contrast, only 8% of the additional deaths occur to working age men and 4% to working age women. This result further emphasizes the likelihood that something other than changes in individuals' own work behaviors drives pro-cyclical mortality fluctuations.

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[5] These numbers are derived by taking the age-group-specific coefficients and multiplying them by the number of deaths occurring in that age group in 2006. Note that the overall estimated effect (-.0029) is slightly smaller than in Table 1 (-.0033). The overall effect given at the bottom of Table 2 is based on a weighted average of the age-specific coefficients, where the weights are the age-specific number of deaths across all years in our sample. The difference between the overall estimates in Tables 1 and 2 arises because the estimate in Table 2 is based on a a less restrictive set of assumptions, using estimates produced by separate regressions for every age group. In particular, we find that allowing different state-specific trends and fixed-effects for each age group generates some differences in the overall estimate. Table 2 is our preferred specification, primarily because of the greater flexibility it allows.

[6] This calculation is slightly complicated by the fact that some age groups have positive (but always statistically insignificant) coefficient estimates associated with the unemployment rate. Thus, we can calculate the total number of additional deaths across narrow age groups as a gross number (in which cells with positive coefficients are ignored) or as a net number, by simply adding up all the positive and negative predicted deaths. In this example, it matters relatively little: 25 to 64 year olds account for 9 percent of all additional deaths using the net number of deaths, and 11 percent using the gross number. In the remaining calculations in this section, we use the net numbers—adding and subtracting across all age categories.

III.B. Mortality Patterns by Cause of Death

Estimating equation (1) separately by cause of death also gives us some hints about the relative role of “internal” vs. ”external” factors. Table 3 provides the results from this exercise[7]. Deaths by motor vehicle accidents are associated with the largest coefficient estimate (-0.025), followed by degenerative brain disease, kidney, respiratory and other accidents (for which each coefficient estimate is equal to -0.009). It is likely that motor vehicle deaths fluctuate because people drive more during strong economic times, but it is not clear that this effect reflects changes in individuals’ own employment status. When we decompose motor vehicle estimates by age, we find that the estimated unemployment coefficients are similar across age groups, consistent with underlying mechanisms that are unrelated to individuals’ own work-related behavior. A similar story applies to the “other accidents” category, for which 63% of additional deaths are among those outside of the prime working ages of 25-65.

Degenerative brain disease contributes the next largest number of deaths. This cause of death seems unlikely to be affected by changes in the own opportunity cost of time or by changes in production processes. Kidney disease is strongly cyclical, but it accounts for a relatively small fraction of additional deaths. Of the three categories, respiratory deaths are the most plausibly related to a trade-off between work and particular health-related activities, but the distribution of these cyclically induced deaths is not consistent with this interpretation. Approximately 80 percent of the averted respiratory deaths are among those over age 60.

The coefficient estimate associated with cardiovascular deaths is not as large as the estimates described above (-0.0036, across all age groups), but because cardiovascular disease is the leading cause of death, this category accounts for many of the remaining deaths. Variation in this cause of death is

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[7] In Appendix 1 we discuss how we classify deaths into the 13 categories considered.

consistent with healthy behaviors that fluctuate with work, so it is tempting to conclude that health behaviors are a significant factor behind the cyclical patterns. However, as Ruhm (2007) has shown, even among cardiac deaths, age patterns are not consistent with mechanisms that rely on a strong connection with individuals' own employment status. Like Ruhm, we find that virtually all of the additional cardiovascular deaths are among those over age 65. Thus, while cardiac deaths contribute a great deal to aggregate mortality fluctuations, the age-specific patterns do not support the hypothesis that most of these additional deaths result from work-related stress or related behavioral changes.[8]

III.C. Unemployment and Employment Indicators by Age and Gender

We continue to investigate the importance of own versus other responses to the business cycle by looking at how mortality rates for different subgroups of the population respond to variation in that subgroup's unemployment rate relative to variation in the unemployment rates corresponding to other groups. If most of the mortality effect is driven by changes in "own" behaviors then the group's own unemployment rate should have the strongest impact on that group's mortality[9]. To investigate this possibility we use the CPS to calculate unemployment and employment rates for broad age/gender groups. We then re-estimate equation (1) separately for men and women who are 25-44 years old, 45-61 years old, and older than 62. For this part of the analysis, we focus on regressions that replace the unemployment rate with the age-group-specific employment-to-population ratio, because the unemployment rate may not capture differences in unemployment among the elderly. This change means that we now expect the estimated coefficient on the regressor of interest to be positive. To facilitate

[8] A few additional coefficient estimates by cause of death are worth noting. Suicide is positively associated with the unemployment rate, as might be expected if job loss affects stress and depression (Ruhm 2000, 2003). Unlike previous studies focusing on US data we also find that cancer deaths are positively correlated with unemployment. These causes of death, however, contribute little to the overall elasticity.

[9] Miller and Paxson (2006) estimate the impact of "relative income" using specifications similar to this, using cross-section and decadal-difference variation in the data.

comparison with earlier results, the first column of Table 4 shows estimated coefficients on the overall unemployment rate. The second column substitutes the overall employment-to-population ratio, and the third column for each group uses the three age-specific employment-to-population ratios. The estimated coefficient on the own-group employment-to-population ratio is shaded in gray.

Table 4 produces little support for the notion that an individual's own group employment rate is driving the cyclicality of mortality. Few of the own group employment coefficients are statistically significant, and most of the estimates on own group employment that are statistically significant are in the opposite direction of the overall finding. Only for women between the ages of 45 and 61 is the own group employment-to-population ratio significantly positively associated with mortality.

Table 4 also shows that mortality fluctuations among the oldest group (over age 62 in this table) are primarily driven by employment changes among younger individuals. This provides an important insight into our earlier finding that pro-cyclical mortality is mainly driven by the top end of the age distribution. It is not this critical group's own employment status that drives the relationship, but rather the employment status of the younger groups.

We have conducted a similar analysis in which we further disaggregate by race (black or white), and then include age- and gender-specific employment rates both for one's own racial group and the other racial group.[10] Once again, we find that own group mortality is, if anything, negatively correlated with one's own group employment rate (defined now by age, gender, and race). In only one case out of 18 is the estimated coefficient on own group employment positive and statistically significant.

III.D. Place of death and health care inputs

[10]Because of concern about small cell sizes, we include only blacks and whites in the by-race analyses.

Since our analyses indicate that mortality fluctuations over the business cycle are concentrated among the elderly, we now explore mechanisms that do not involve changes in individuals' own employment status or time use. For example, there may be cyclical changes in the quality, quantity or nature of health care inputs, which are relatively heavily utilized by those over age 65. There is evidence that employment in the health care sector is lower during expansions than during recessions: Goodman (2006) finds that the correlation between changes in hospital employment and changes in aggregate employment is strongly negative (-0.90) and concludes that "… at times of peak U.S. hiring, when the labor shortage in hospitals may be particularly intense, hospitals with staffing shortages may face restrictions on the volume of business that can be performed at a particular time."

Another important clue to the mechanism at work here involves common claims that nursing homes experience especially severe shortages of nursing aides when the economy is strong. Yamada (2002) notes that during the late 1990's very low unemployment rates exacerbated already severe labor shortages for direct care workers—nursing aides, home health workers, and other paraprofessional caregivers. She cites a study from the state of New York that suggests that between 70 to 90% of home health care agencies and nursing homes indicate shortages of direct care workers. If such shortages become particularly acute during good economic times, then we might expect higher mortality among the elderly to follow.[11]

This possibility is reinforced by our finding that cyclicality among older women is greater than among older men. Older men and women tend towards different living arrangements, and may therefore be differentially affected by fluctuations in the quality of health care. For example, because women tend to be married to older men, and because men have a shorter life expectancy than women, older women are

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[11] Yamada also refers to a long list of studies that point to macroeconomic conditions as an important factor driving these labor shortages.

relatively more likely to have a market-based caregiver than are older men, and they are more likely to reside in nursing homes at the end of their lives (Murtaugh, et al. 1990).

In this section, we investigate whether changes in health care inputs over the business cycle might be part of the story. We explore this possibility by comparing mortality patterns in nursing homes to mortality patterns elsewhere, and by comparing estimated business cycle effects across states where the elderly have different types of living arrangements. We also look at how occupations that are typically associated with nursing home care vary with fluctuations in the economy.

III.D.1. Direct Evidence from Vital Statistics Place of Death data

We begin by comparing patterns of mortality among individuals living in nursing homes to the patterns among those living alone or with family members. Our focus on nursing home residents is motivated by the fact that this is a group certain to be in contact with health care workers near the time of death, and, as such, more likely to be affected by changes in the quality of care that is available over the business cycle. In addition, the evidence discussed above suggests that nursing homes are particularly vulnerable to the types of labor shortages that might reduce the quality of care that is provided during an economic boom.

The starting point of our investigation is the place of death information that is provided by the Vital Statistics mortality files after 1983. Death certificates indicate whether the death occurred in a hospital, nursing home, residence, or other location. Place of death is only a weak proxy for where an individual was living prior to death because, for example, many nursing home residents will die in a hospital after being transferred there for an illness. However, information on place of death is readily available and so a natural place to start.

Unfortunately, the Vital Statistics place of death codes are available only after 1982 and, over the period we analyze, they have experienced two substantial changes. First, in 1989 death certificates were changed such that physicians no longer filled out an open-ended question regarding the deceased individual's place of death. Instead, they began to fill out boxes indicating whether the death occurred in a hospital, residence or nursing home. For years prior to 1989, the categories listed in the Vital Statistics codebook include hospitals (and several subsets), "other institutions providing patient care", and all other reported places. We assume that "other institutions providing care" are primarily nursing homes. Second, in 2003 the categories associated with the boxes were changed slightly. "Nursing home" was replaced by "nursing home/long-term care" and a separate category was added for deaths in a hospice. We have recoded the categories across years into: nursing homes, hospitals, and "other." Figure 3 shows the fraction of deaths occurring in each category by year and suggests that these changes did not have a major impact. The timing of the coding changes is not associated with any dramatic breaks in any of the series.

Table 5 presents results based on data from the years 1983 -2006 and 1983-2002. Our dependent variable is age-adjusted mortality among those 65 and over.[12] For the longer time period, the estimated impact of the unemployment rate on elderly mortality (-0.002) is similar to our main results. When deaths are divided between nursing homes and non-nursing homes, the estimated coefficient is not statistically significant for either sub-group, although the estimated coefficient for deaths that occur in nursing homes is very large (-.03) and negative, which suggests that nursing home deaths may play an important role. Dropping deaths in nursing homes and focusing on all other deaths (nearly 80 percent of all deaths) changes the coefficient on the unemployment rate from negative and significant to positive and

[12] We include the second set of regressions in order to show our estimates' sensitivity to the change in place of death coding starting in 2003. We have also restricted the sample to the years 1989 through 2002 (the period for which there are no changes in place of death coding). For the shorter period 1989 to 2002, however, we find no statistically significant relationship between the unemployment rate and overall mortality, or the unemployment rate and mortality by place of death.

not statistically significant. Mortality in nursing homes thus seems to be a very important part of overall cyclicality, which is, again, suggestive of mechanisms that have little to do with work, time use or health behaviors.

In the next panel of Table 5 we drop the years after 2002, when the second change in the coding of place of death took place. As noted in Section II, dropping the last four years of observations substantially increases the magnitude of the estimated coefficient for the full sample. The estimate continues to be driven by deaths in nursing homes, where we observe a statistically significant coefficient estimate of -.056. In contrast, among deaths taking place elsewhere, the estimated impact of the unemployment rate is positive, at 0.005 (but not statistically significant). This pattern holds for both men and women. Focusing on this shortened period gives relatively more weight to the late 1990s, when unemployment rates were particularly low and labor shortages within low-skilled health care occupations may have been particularly acute. Nevertheless, in both sample periods, nursing home deaths are associated with an estimated coefficient that is *an order of magnitude* larger than the coefficient that is estimated among deaths taking place elsewhere.

III.D.2. Combining information on place, cause, and age of death

We next use the Vital Statistics data to decompose the overall cyclical coefficient by age, cause and place of death. Category specific coefficient estimates are displayed on the left side of Table 6. The overall coefficient (such as reported in Table 1) is approximately equal to the weighted average of coefficients from each category, where the weights are the number of deaths in the category (shown on the right side of Table 6). Multiplying each coefficient estimate by the corresponding number of deaths and then computing a weighted average produces an overall coefficient estimate of -.0038.

The category-specific semi-elasticity estimate can be used to calculate counterfactual values of the overall level of cyclicality that would exist if certain age/cause/place categories were not cyclically sensitive. For example, motor vehicle accidents represent only around 2% of total deaths in a given year, and so account for a relatively small part of the relationship between the economy and mortality.[13] If the cyclical coefficients on motor vehicle deaths were all equal to zero, then the semi-elasticity would fall only slightly, from -.0038 to -.0033. In contrast, nursing home deaths among those over age 65 account for approximately 18% of all deaths (410,136/2,265,609). If the estimated coefficient associated with deaths in nursing homes is set to zero, then the estimated impact of the unemployment rate becomes positive (0.002). Viewed another way, allowing *only* for the cyclicality occurring in nursing homes, and setting all non-nursing home death coefficients to zero, the overall cyclicality term would be equal to -.0058. Nursing home deaths among those over 65 more than account for total cyclical mortality. This "over-explanation" happens because the estimated coefficients for non-nursing home deaths are slightly positive. Thus, while it is certainly true that by-cause breakdowns of non-elderly, non-motor vehicle deaths uncover categories that move pro-cyclically, their effect relative to total cyclical mortality is small.

III.D.3. Interactions with institutionalized fraction of the Elderly

Place of death is a weak proxy for where an individual was living prior to his death. We therefore supplement our Vital Statistics analyses with a different approach that makes use of information on residence in institutional group quarters available in the Census. The Census does not report residence in nursing homes specifically, but it does indicate whether individuals are living in group quarters, which typically include military barracks, nursing homes, college dormitories and prisons. For individuals over

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[13] Motor vehicle accidents, given their large coefficient do account for a substantial fraction (around 12%) of all cyclically induced deaths.

age 65 it is likely that the vast majority of residents in group quarters are living in nursing homes. Looking at the fraction of older individuals who live in group quarters in each state gives a reasonable approximation of which states have relatively high or low fractions of their elderly population in nursing homes. If nursing home staffing or quality of care is an important component of cyclical mortality we should expect to see greater cyclicality in states with larger nursing home populations.

To investigate this possibility, we begin by using data from the 1980, 1990, and 2000 Census files to calculate the fraction of individuals age 65 and over living in group quarters in each state and Census year. Table 7 summarizes the fraction of different age groups living in group quarters and the dispersion in this measure across states. As has been documented elsewhere, the fraction of elderly living in nursing homes has declined since 1980, as a result of both improved health among older Americans and the growth of non-institutional, "assisted living" residences. Between 2 and 4 percent of men over age 65 lived in a nursing home during our sample period; whereas 5 to 7 percent of elderly women resided in a nursing home. The lower panel of Table 7 shows the extent of variation across states in these fractions. Going from the 25^{th} to the 75^{th} percentile of the distribution of (state-level) nursing home residence moves the fraction by one to two percentage points for both men and women.

In the top half of Table 8 we show what happens when we add to our main regression specification (equation 1) an interaction between the unemployment rate and the state's fraction of women (or men) over age 65 who are living in nursing homes. The dependent variable in these regressions is the log of the age-adjusted mortality rate for women (or men) over age 65. In the lower half of the table (as a placebo test), the dependent variable is the log of the age-adjusted mortality rate for ages 0-45. For simplicity, we present results using fraction of the elderly living in group quarters from the 1980 Census only, so that they do not vary with time. For women, the interaction between the unemployment rate and the fraction in nursing homes is negative and statistically significant, suggesting that the pro-cyclicality of mortality is stronger in states where larger fractions of the elderly reside in nursing homes. The estimate suggests that increasing the unemployment rate by 1 percentage point would decrease mortality rates by 0.4 % in a state at the 25th percentile of the distribution of the fraction in nursing homes and by 0.5% in a state at the 75th percentile.[14] For men, the estimated coefficient on the interaction term is smaller and not statistically significant. The estimated magnitudes for women are consistent with the estimates produced by our analyses using place of death from the mortality files. If we calculate an out-of-sample prediction (using the results from Table 8) of the implied coefficient on the unemployment rate for an area in which *all* elderly residents live in nursing homes, we predict that the effect of a one-percent increase in the unemployment rate is approximately -.03 to -.06. This is similar in magnitude to the estimates shown in Table 5.

In the lower half of the table we perform a placebo test in which we repeat this exercise, using age-adjusted deaths among those ages 0 to 45 as the dependent variable. If the estimates in the top half of the table are capturing something about changes in the quality of nursing home care (and not some unobserved feature of the state) then the interaction terms in the lower half of the table, which relate mortality at younger ages to the fraction of elderly in nursing homes, should not be statistically significant. In fact, neither interaction is close to statistical significance, although the large accompanying standard error estimates also make it impossible to rule out effects that are similar in size to the estimates for those 65 and over.

In sum, both our Vital Statistics and Census analyses indicate that pro-cyclical mortality cannot be explained without focusing on the elderly, particularly elderly persons who reside in nursing homes. We have attempted to further explore these relationships using individual level data from the HRS and

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[14] This interaction also explains the difference between elderly men and women from the bottom of Table 1. If we use the estimates in Table 8, evaluated at the median fraction in nursing homes for men and women (from Table 7) we get a total effect of the unemployment rate for women of -.0046, and for men of -.0020.

AHEAD datasets. Unfortunately, these datasets have important limitations that limit their usefulness for our analysis, and which lead to inconclusive results. We describe the analyses and results in Appendix 2.

Overall, we view our findings on the importance of mortality in nursing homes as an important clue to the mechanisms driving pro-cyclical mortality. This does, however, raise the question of why elderly nursing home mortality exhibits such a strong pro-cyclical pattern. One likely possibility, suggested by earlier work noted above, is that there is variation in the quality of health care that is provided over the business cycle; we explore this possibility in the next section.

III.D.4. Evidence from institution-level measures of health care labor inputs

Why are mortality rates among nursing home residents particularly sensitive to the local economy? One possibility is that the quality of care provided in nursing homes falls when labor markets are tight. We are unable to look directly at how the quality of nursing home staff changes with the business cycle, but data from the Online Survey Certification and Reporting Database (OSCAR) allow us to estimate how the number of hospital and nursing home workers changes with the unemployment rate. OSCAR includes data on any institutional healthcare provider that is certified to provide services under Medicaid or Medicare. The dataset covers 97% of all hospital facilities in the US, and contains detailed information on staffing levels. OSCAR's standard analytical files are available from 1991 through 2007. From 1992 forward, OSCAR's Hospital Service Area file includes information on total patient caseloads and days of care, for several types of facilities, including skilled nursing facilities.

We use these data to examine changes in skilled nursing facility staffing levels over the business cycle. Table 9 shows estimates of the relationship between the log of employment (or occupation-specific employment, including physicians, registered nurses and licensed practical nurses, certified aides, and "other") in skilled nursing facilities and the unemployment rate. The estimates are produced by regressions that include either state-level fixed effects or provider-level fixed effects, along with state-specific trends. The regressions are weighted by the provider size, or by the total number of beds.[15]

We find that staffing levels in nursing homes rise during periods of high unemployment.[16] A one percentage point increase in the unemployment rate raises total full-time employment at skilled nursing facilities by approximately three percent. There is no statistically significant increase in the number of physicians, but there are significant increases in nurses, certified aides, and other occupations. Because physicians are an extremely small part of total employment in nursing homes, the non-MD categories all rise by approximately the same amount as total employment, around 3% for a one percentage point increase in the unemployment rate. While these estimated effects are statistically significant, their substantive importance is not immediately clear. Consider a four percentage point change in the unemployment rate, which is a reasonable approximation of the magnitude of the change that accompanies a boom to bust. Such a change would imply an increase in staffing levels of approximately 12 percent. There are few persuasive studies connecting lower-skilled staffing levels to quality of care outcomes, but our earlier results suggest that such a change in unemployment would reduce annual mortality rates in nursing homes by approximately 12 percent.

To supplement our OSCAR analysis we have also used the CPS to construct measures of employment, and employee characteristics, for several health care occupations. Specifically, we construct measures of the fraction of the population in each state-year cell who are MDs, Nurses (which can be further divided into registered nurses (RNs) and licensed practical nurses (LPNs), and Aides (also divided

[15] Unfortunately, the files with staffing levels in skilled nursing facilities do not contain patient counts; patient count levels are provided only for the hospital files.

[16] We have conducted a similar exercise for hospitals, but find little evidence that changes in hospital staffing levels are related to the business cycle.

into Health Aides and Nursing Aides).[17] For occupations other than MD, where there may be some flexibility in the level of education required, we also calculate the fraction of each occupation with particular levels of education. Unfortunately, in 2003 there are changes in the occupational coding of some of the key health care occupation categories. As a result, this analysis only includes the years 1983 through 2002.

We complement our OSCAR analyses by using these measures to examine how these occupations fluctuate with the business cycle. Specifically, we regress these occupation specific fractions on a set of state and year fixed effects, state-specific trends, and the same set of state-year demographic controls that are included in our main mortality regressions. The results of this exercise are summarized in Table 10, and confirm that employment in low-skilled health occupations moves counter-cyclically. Each row of Table 10 shows results from a separate regression in which the dependent variable is the fraction of a state's employment in a given occupation. This dependent variable is different from that used in our OSCAR analyses because it captures employment in the given occupation across all types of employers (not only nursing homes). Also unlike our OSCAR analyses, occupational categorization in the CPS is determined by workers rather than employers. The CPS regressions are also based on a different set of years (1983-2002 instead of 1990-2006).

The first rows of Table 10 show results for the fraction employed as doctors and nurses. The estimates provide evidence of pro-cyclicality among these more skilled occupations. In the lower half of the table are the results for health and nursing aides, and then all aides (the sum of the previous two rows). We find evidence that aides exhibit counter-cyclical patterns; for all aide categories, a higher unemployment rate is associated with a statistically significant increase in employment. A one percentage point increase in the unemployment rate raises the fraction of employment in aide occupations

[17] Specifically, our category of nursing aides corresponds to CPS occupation 447 (prior to 2003), "nursing aides, orderlies, and attendants" and health aides corresponds to CPS occupation 446 "health aides, except nursing."

by more than 1 percent. Our finding that aides exhibit countercyclical employment is consistent with our earlier evidence on nursing home deaths, since nursing aides are heavily concentrated in skilled nursing facilities (GAO, 2001).[18]

IV. Conclusion

In this study we confirm a robust link between mortality and unemployment rates. Findings by Ruhm (2000, 2003, 2005) are robust to a number of changes in the underlying data, additional controls, and to including an additional decade of data. We show that adjusting mortality rates for changes in the age distribution can be quite important, and that when using recent data this appears to increase the magnitude of the estimated coefficient of unemployment on mortality.

Our primary contribution is to bring additional data to bear on the question of why such a relationship exists. Specifically, we focus on age- and cause-specific patterns of the cyclicality of mortality rates. We confirm that the largest responses (in terms of estimated coefficients) are among relatively young adults, who also have the most cyclically sensitive employment responses. However, children have responses that are just as large, casting doubt on mechanisms that rely primarily on changes in own work hours or employment status. Furthermore, mortality fluctuations among the non-elderly are mostly driven by the cyclicality of motor vehicle accidents, which is roughly similar in magnitude across both working and non-working aged individuals. We also show that the overall association of unemployment and mortality is driven by the mortality response among those over 65, particularly women over age 65.

Our analysis next compared the responsiveness of mortality within demographic groups (defined by age, sex, and race) to their group-specific measures of employment. Own-group labor market indicators are not positively related to that group's mortality, and there is some evidence that the

[18] The GAO reports that, in 1999, nearly half of all nursing aides worked in nursing homes, with the remainder split relatively evenly between hospitals and home health care agencies.

relationship may be negative. These two sets of findings lead us to conclude that mechanisms that do not involve individuals' own employment status or time use over the business cycle may play an important role.

Given the apparent importance both of mortality after age 65 and of mechanisms not involving individual employment changes, we next focus on mortality at older ages and the potential role for changes in health care inputs. Three sets of empirical findings support an important role for health care inputs among the elderly as an important factor explaining pro-cyclical mortality. First, using the place of death recorded on the death certificates, we show that deaths occurring in nursing homes are particularly responsive to the state unemployment rate. Second, we use Census data to calculate the fraction of each state's elderly population residing in nursing homes, and show that mortality in states with higher fractions of nursing home residents is more pro-cyclical. Decompositions of the overall cyclicality of mortality point strongly to these nursing home deaths as a key mechanism. Finally, we show that employment levels in skilled nursing facilities show statistically significant declines when the unemployment rate falls, findings that are corroborated by occupation data from the CPS. Taken together, these findings suggest that pro-cyclical mortality may involve mechanisms that have little to do with individual behavioral changes in employment or time use over the business cycle. Further, our findings provide new evidence that staffing difficulties among relatively low-skilled nursing occupations may be an important focus for efforts to improve the quality of health care.

Appendix 1 – Mortality classifications

We grouped causes of death into 13 categories according to codes and recodes published by the Center for Disease Control for each version of the International Classification of Diseases (ICD). More about the ICD can be found at http://en.wikipedia.org/wiki/International Statistical Classification of Diseases and Related Health Problems. We used the ICD-8, ICD-9 and ICD-10, which correspond to the years 1968-1978, 1979-1998 and 1999-2004. The ICD groups reported causes of death into categories called "recodes." Within a given year, there are several different sets of re-codes. For example, the 1990 data include a 272 cause recode, a 72 cause recode and a 61-cause recode. For deaths occurring in 1968-1978 we use the ICD-8 69-cause recode. We use the ICD-9 72-cause recode for deaths occurring in 1979-1998, and the ICD-10 113-cause recode for deaths occurring in 1999-2004. Column 2 of Appendix Table 1 shows how we assigned each re-code category into our 13 causes of death groups.

There are some causes of death which the CDC recodes as "other and unspecified" but which seemed to fit our categories. For example, in the ICD8 years, code 782.4 was classified as "other" by the 69-cause recode. Further investigation of this recode, however, made clear that this "other" category was related to cardiovascular deaths, so we reclassified all deaths that had been assigned the 782.4 code as cardiovascular deaths. The third column of Appendix Table 1 shows which "other" recodes in the ICD were reclassified into each of our 13 causes of death groups.

Appendix Table 1

Classification of ICD Recodes into Cause of Death Categories

	Cause of Death	Recodes	Raw codes reclassified from "other
1	Cardiovascular	ICD8: 310,320,330,350,360,370,380,390,400,410,430,440,450,460,470,480 ICD9: 320,330,340,360,370,380,390,400,410,420,440,450,460,470,480,490 ICD10: 55,56,57,59,60,62,63,65,66,67,68,69,70,71,73,74,75	ICD8: 7824
2	Cancer	ICD8:150,160,170,180,190,200,210,220,230,240 ICD9: 160,170,180,190,200,210,220,230,240,250 ICD10: 20, 21, 22, 23, 24, 25, 26, 27, 28, 29, 30, 31, 32, 33, 34, 35, 36, 38, 39, 40, 41, 42, 43, 44	
3	Respiratory	ICD8:490,510,520,540,550,560 ICD9: 500,520,530,550,560,570,580 ICD10: 77,78,80,81,83,84,85,86,87,88,89	ICD8: 450, 510, 5112, 512, 513, 514, 5150, 5151, 5159, 517, 518, 5190, 5191, 5192, 5193, 5199 ICD9: 500, 501, 5070, 5109, 5119, 512, 5130, 514, 515, 5168, 5183, 5184, 5185, 5188, 5191, 5198, 7991 ICD10: R092
4	Infections and immune deficiency	ICD8:10,20,40,50,60,70,80,90,100,110,120,130,280 ICD9: 10,20,40,50,60,70,80,90,100,110,120,130,140,290 ICD10: 1,2,3,5,6,7,8,9,10,11,13,14,15,16,17,18,50	ICD8:5990 ICD9: 2791,2793,5990 ICD10: D849,N390
5	Degenerative brain diseases	ICD10: 51,52	ICD8: 2900, 2901, 299, 3049, 3099, 323, 3304, 340, 341, 342, 3439, 3441, 3449, 3451, 3452, 3459, 3471, 3479, 794, 7330 ICD9: 2900, 2901, 2902, 2904, 2949, 2989, 3109, 311, 319, 3239, 3310, 3314, 3319, 3320, 3352, 340, 3419, 7855, 797, 3440, 3451, 3453, 3459, 3481, 3483, 3485, 3489, 3568, 3580 ICD10: F019, F03, F069, F109, F179, F329, F79, G122, G319, G35, G409, G419, G459, G700, G809, G825, G919, R54, R628, G934, M349
6	Kidney and urethra	ICD8:630,640,650 ICD9: 650,660,670,680 ICD10: 98,99,100,101,102	ICD8:465,592,5931,5932,595,5999,792 ICD9: 5920,5939,5996 N289
7	Nutrition-Related	ICD8:250,260,270 ICD9: 260,270,280 ICD10: 45,46,48,49	ICD8:277 ICD9: 2720,2724,2780 ICD10: E668,E669,E780,E785
8	Motor Vehicle Accidents	ICD8:770 ICD9: 800 ICD10: 114	
9	Other Accidents	ICD8:780 ICD9: 810 ICD10: 115,116,118,119,120,121,122,123	
10	Suicides	ICD8:790 ICD9: 820 ICD10: 125,126	
11	Homicides	ICD8:800 ICD9: 830 ICD10: 128,129,130	
12	Other, Unspecified and ill-defined	ICD8:Remainder of 740,750 ICD9: Remainder of 770,780 ICD10: Remainder of 110,111	
13	Misc: Birth defects; diseases of blood/ bone/ gastrointestinal/ metabolic/ autoimmune; drug abuse; and other external	ICD8:570,580,590,600,610,660,680,690,700,720,730,810 ICD9: 590,600,610,620,630,690,710,720,730,750,760,840 ICD10: 90,91,92,94,95,96,103,104,106,107,108,109,132,133,134,135	ICD8: 251, 2730, 274, 276, 279, 2869, 2871, 288, 2910, 3032, 3039, 3310, 4510, 4519, 453, 4549, 4560, 4589, 5301, 5304, 5309, 5340, 535, 5379, 5621, 5630, 5631, 5670, 5679, 5693, 5694, 5699, 570, 5730, 5739, 5760, 5769, 5770, 7123, 7130, 715 ICD9: 2733, 2762, 2765, 2769, 2770, 2773, 2779, 2866, 2869, 2873, 2875, 2880, 2898, 303, 3049, 3050, 3059, 3334, 3591, 4512, 4519, 4560, 4589, 4590, 4599, 5301, 5303, 5305, 5559, 556, 5570, 5571, 5579, 558, 5621, 5672, 5679, 5698, 570, 5722, 5724, 5728, 5733, 5739, 5761, 5770, 5771, 5789, 7140, 7330, 7373 ICD10: D65, D689, D696, D70, D869, E854, E859, E86, E870, E872, E875, E878, E889, F101, F102, F141, F191, F192, G10, K222, K224, K509, K529, K550, K559, K562, K566, K572, K578, K579, K631, K639, K650, K659, K729, K759, K760, K767, K830, K85, K860, K861, K922, M069, M321, M819, M869

Appendix 2 – HRS / AHEAD

In this appendix we discuss our analyses that combine individual-level residence in nursing homes and mortality (or other health outcomes) using data from the Health and Retirement Study (HRS) and Asset and Health Dynamics of the Oldest Old (AHEAD). The AHEAD data include cohorts of individuals born before 1923, and begins to follow them around age 70. The HRS sample includes cohorts born between 1931 and 1941 and follows them beginning around age 50. Reliable mortality information for HRS and AHEAD participants who cease responding to the questionnaire is verified by the HRS staff through a match to the national death index. Unfortunately, given what we have found about the importance of nursing home deaths, the HRS/AHEAD data have an important limitation: the initial sampling frame included only non-institutionalized individuals. The surveys do follow individuals as they transition into nursing home residency, however. So, in later years of the sample, some of the respondents are residing in nursing homes.

Using the HRS/AHEAD data for 1992-2006, we have estimated both individual-level mortality hazard models and state-year cell level specifications that are similar to our main specification. We find no evidence of that mortality among this population is related to state-level unemployment rates. This could reflect either reduced statistical power given the smaller sample sizes and shorter time-series relative to the other datasets that we use, or it could reflect the underrepresentation of the nursing home population in these data. We have also attempted to look at the cyclicality of mortality among those respondents that transitioned into nursing homes in the later years of the HRS, but we are hampered by the fact that our samples of nursing home residents are so small. Approximately 2400 individuals are observed to transition into nursing homes over the sample period, and so our effective state-year cells are very small. We do not find any evidence that mortality is statistically significantly related to the unemployment rate among HRS/AHEAD nursing home residents, although point estimates of the coefficient on the interaction term between the unemployment rate and nursing home residency are consistently negative.

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Table 1

Replication of Ruhm (2000) estimates, different specifications and years

year	Ruhm- all 1972-1991	Ruhm - all 78-91	Ruhm rhs Our mortality 78-91	Our rhs Our mortality 78-91	Our rhs Our mortality 78-06	Age-adjusted mortality 78-06
weight by pop						
coeff	-0.0054	-0.0048	-0.0038	-0.004	-0.0019	-0.0033
se	(0.0010)	(0.0010)	(0.0007)	(0.0008)	(0.0010)	(0.0010)
no weights						
coeff	-0.0056	-0.0053	-0.004	-0.0052	-0.0028	-0.0037
se	(0.0009)	(0.0010)	(0.0008)	(0.0010)	(0.0008)	(0.0007)
weight by pop						
women						
coeff				-0.0034	-0.0023	-0.004
se				(0.0009)	(0.0009)	(0.0011)
men						
coeff				-0.044	-0.0013	-0.0024
se				(0.0009)	(0.0011)	(0.0011)
women 65+						
coeff				-0.0043	-0.0023	-0.0041
se				(0.0009)	(0.0010)	(0.0012)
men 65+						
coeff				-0.0035	-0.0003	-0.0018
se				(0.0009)	(0.0008)	(0.0006)

Parameters are estimated Mortality semi-elasticity with respect to the state-year unemployment rate. Controls include State and Year Fixed effects, State-specific trends, demographic and education controls. Standard errors clustered at the state level. Weighed estimates use state-year population of relevant group. "Ruhm rhs" has control variables as in Ruhm (2000), "Our rhs" has control variables as described in text.

Table 2
Effects of Unemployment on Mortality by Age Coefficients and Additional Deaths

	Coefficients			Additional Deaths from 1% Increase in Unemployment Rate (based on 2006 #s of deaths)		
Age Group	All	Men	Women	All	Men	Women
0 to 4	-0.014	-0.015	-0.012	-464	-278	-176
	(0.003)	(0.003)	(0.003)			
5 to 9	-0.009	-0.01	-0.008	-25	-16	-9
	(0.005)	(0.005)	(0.009)			
10 to 14	-0.005	0.003	-0.018	-17	6	-24
	(0.005)	(0.004)	(0.009)			
15 to 19	-0.014	-0.015	-0.011	-192	-149	-42
	(0.004)	(0.005)	(0.005)			
20 to 24	-0.018	-0.018	-0.015	-381	-291	-75
	(0.004)	(0.004)	(0.004)			
25 to 29	-0.008	-0.011	0.002	-167	-167	11
	(0.004)	(0.005)	(0.004)			
30 to 34	-0.004	-0.008	0.004	-88	-120	28
	(0.006)	(0.007)	(0.006)			
35 to 39	0.001	0.000	0.004	31	0	45
	(0.005)	(0.007)	(0.004)			
40 to 44	0.002	0.001	0.003	103	32	59
	(0.004)	(0.005)	(0.003)			
45 to 49	0.000	0.001	-0.002	0	49	-61
	(0.002)	(0.003)	(0.003)			
50 to 54	-0.002	-0.001	-0.003	-212	-66	-119
	(0.002)	(0.002)	(0.002)			
55 to 59	-0.001	0.001	-0.002	-133	81	-103
	(0.002)	(0.002)	(0.002)			
60 to 64	-0.001	-0.001	-0.001	-148	-88	-61
	(0.001)	(0.001)	(0.001)			
65 to 69	-0.002	-0.002	-0.002	-344	-197	-146
	(0.001)	(0.001)	(0.001)			
70 to 74	-0.004	-0.003	-0.005	-873	-358	-494
	(0.001)	(0.001)	(0.001)			
75 to 79	-0.002	-0.001	-0.003	-595	-154	-430
	(0.001)	(0.003)	(0.002)			
80 to 84	-0.003	-0.003	-0.004	-1109	-519	-786
	(0.001)	(0.001)	(0.001)			
85+	-0.003	0.000	-0.004	-2106	0	-1842
	(0.002)	(0.002)	(0.002)			
Sum of deaths				-6720	-2234	-4224
wtd avg	-0.0029	-0.0022	-0.0035			

First 3 columns ("coefficients") report mortality semi-elasticity with respect to the state-year unemployment rate. Controls include State and Year Fixed effects, State-specific trends, demographic and education controls. Standard errors clustered at the state level. Weighted by population. Last row weighted by age-group-number of deaths over the 1978-2006 period. Last 3 rows compute "averted deaths" by multiplying the coefficient by the # of 2006 deaths to each age group.

Table 3

Effects of Unemployment on Mortality by Age & Cause and Additional Deaths

	Coefficients on Unemployment Rate Age Group				Additional Deaths from 1% Increase in Unemployment Rate Age Group			
Cause of Death	All	0 to 24	25-64	65 and up	All	0 to 24	25-64	65 and up
cardiovascular	-0.0036	-0.0026	-0.0007	-0.0031	-2234	-7	-106	-2115
cancer	0.0025	-0.0001	0.0021	0.0033	1660	0	355	1315
respiratory	-0.0087	-0.0210	-0.0091	-0.0060	-1497	-37	-277	-1173
infections & immune deficiencies	-0.0094	-0.0111	-0.0181	-0.0091	-1019	-16	-517	-478
degenerative brain diseases	-0.0094	0.0028	-0.0076	-0.0099	-1933	3	-91	-1836
kidney	-0.0093	0.0053	-0.0069	-0.0087	-394	2	-56	-331
nutrition related	-0.0013	-0.0016	-0.0024	-0.0007	-115	-1	-62	-44
MVA	-0.0254	-0.0269	-0.0247	-0.0194	-1120	-351	-622	-137
other accidents	-0.0089	-0.0156	-0.0008	-0.0094	-447	-128	-31	-278
suicide	0.0138	-0.0022	0.0227	0.0094	569	-10	536	50
homicide	-0.0094	-0.0145	-0.0066	-0.0034	-186	-100	-74	-3
VS "other"	-0.0188	-0.0482	-0.0281	-0.0046	-1086	-319	-572	-188
Residual	-0.0057	-0.0052	-0.0084	-0.0019	-657	-124	-404	-119

First 3 columns ("coefficients") report mortality semi-elasticity with respect to the state-year unemployment rate. Controls include State and Year Fixed effects, State-specific trends, demographic and education controls. Standard errors clustered at the state level. Weighted by population. Last 3 columns compute "averted deaths" by multiplying the coefficient by the # of 2006 deaths to each age group.

Table 4

Response of Mortality by Age and Gender to Group-Specific Employment-Population Rates

All Sexes		Ages 25 to 44			Ages 45 to 61			Ages 62 and over	
Emp/Pop Ages 25-44			-0.0052**			-0.0012			0.0008*
			(0.0024)			(0.0008)			(0.0004)
Emp/Pop Ages 45-61			0.0047***			0.0006			0.0008**
			(0.0017)			(0.0006)			(0.0004)
Emp/Pop Ages 62+			0.0019			0.0008			0.0001
			(0.0014)			(0.0005)			(0.0003)
CPS Overall Emp/Pop		0.0013			-0.0005			0.0013**	
		(0.0025)			(0.0009)			(0.0006)	
BLS Unemployment	-0.0016			-0.0004			-0.0031***		
	(-0.0042)			(0.0014)			-0.0008		
Men									
Emp/Pop Ages 25-44			-0.0003			-0.0011			0.0010***
			(0.0023)			(0.0010)			(0.0004)
Emp/Pop Ages 45-61			0.0017			-0.0001			0.0002
			(0.0024)			(0.0006)			(0.0003)
Emp/Pop Ages 62+			0.0005			0.0003			-0.0001
			(0.0013)			(0.0005)			(0.0002)
CPS Overall Emp/Pop		0.0026			-0.0019*			0.0007*	
		(0.0031)			(0.0011)			(0.0004)	
BLS Unemployment	-0.0036			0.0008			-0.0020***		
	(-0.0051)			(0.0017)			-0.0006		
Women									
Emp/Pop Ages 25-44			-0.0029*			0.0001			0.0001
			(0.0016)			(0.0006)			(0.0004)
Emp/Pop Ages 45-61			0.0013			0.0010**			0.0009**
			(0.0011)			(0.0005)			(0.0004)
Emp/Pop Ages 62+			-0.0001			0.0003			0.0003
			(0.0013)			(0.0007)			(0.0004)
CPS Overall Emp/Pop		-0.0017			0.0016**			0.0016*	
		(0.0021)			(0.0007)			(0.0008)	
BLS Unemployment	0.0032			-0.002			-0.0041***		
	(-0.0032)			(0.0012)			-0.0012		

N = 1479. See Table 1 for specification. Shaded coefficients represent own-group employment/population.

Table 5

Effect of Unemployment Rate on Log Mortality by Place of Death: Nursing Homes vs. All Other

	1983-2006			1983-2002		
	All	Men	Women	All	Men	Women
All places	-0.002	-0.002	-0.002	-0.004	-0.003	-0.005
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Nursing Homes	-0.028	-0.028	-0.029	-0.056	-0.058	-0.055
	(0.019)	(0.019)	(0.019)	(0.027)	(0.027)	(0.027)
Not Nursing Homes	0.003	0.001	0.005	0.005	0.004	0.006
	(0.003)	(0.002)	(0.003)	(0.003)	(0.003)	(0.004)
N=	1224	1224	1224	1020	1020	1020

Parameters are estimated Mortality semi-elasticity with respect to the state-year unemployment rate. The second and third rows count only deaths in Nursing Homes and all other places--"Not Nursing Homes". Controls include State and Year Fixed effects, State-specific trends, demographic and education controls. Standard errors clustered at the state level. Estimates weighed by population.

Table 6

Unemployment Rate Coefficients by Broad Cause & Place of Death

1983-2006	Cause/Place of Death			Average deaths per year-1983--2006		
Age group	MVA	non-MVA		MVA	non-MVA	
0-23	-0.031	-0.011		14,798	67,362	
	(0.006)	(0.003)				
24-45	-0.029	0.009		14,669	120,317	
	(0.006)	(0.005)				
45-64	-0.018	0.001		8,051	390,768	
	(0.007)	(0.001)				
		Nursing Homes	Not Nursing Homes		Nursing Homes	Not Nursing Homes
65 plus	-0.017	-0.032	0.004	7,216	410,136	1,232,292
	(0.006)	(0.020)	(0.003)			

1983-2002	Cause/Place of Death			Average deaths per year-1983--2002		
Age group	MVA	non-MVA		MVA	non-MVA	
0-23	-0.029	-0.011		15,108	68,749	
	(0.005)	(0.003)				
24-45	-0.03	0.001		14,828	121,671	
	(0.009)	(0.005)				
45-64	-0.023	0.000		7,543	380,693	
	(0.009)	(0.001)				
		Nursing Homes	Not Nursing Homes		Nursing Homes	Not Nursing Homes
65 plus	-0.019	-0.059	0.006	7,190	391,367	1,225,633
	(0.008)	(0.027)	(0.003)			

1224 observations for 1983-2006, 1020 observations for 1983-2002. "MVA" means "Motor Vehicle Accidents."

Table 7

Fraction of State Residents 65 and Over Residing in Institutional Group Quarters

	1980	1990	2000
Means			
(weighted by state population)			
men	0.0363	0.0263	0.0317
	(0.011)	(0.005)	(0.008)
women	0.0635	0.0511	0.0602
	(0.016)	(0.009)	(0.014)
Means			
(unweighted)			
men	0.0385	0.0277	0.0339
	(0.012)	(0.009)	(0.013)
women	0.0651	0.0503	0.0612
	(0.018)	(0.011)	(0.017)
Distribution across states			
(unweighted)			
men			
25th percentile	0.032	0.023	0.025
median	0.037	0.027	0.033
75th percentile	0.04	0.028	0.038
women			
25th percentile	0.059	0.048	0.047
median	0.065	0.049	0.062
75th percentile	0.07	0.055	0.07

Tabulations from Public use Census Files, 1980, 1990, 2000 of fraction of state population aged 65 and over residing in instiutional group quarters. Standard deviations (at state level) in parentheses.

Table 8

Cyclicality of Mortality by Fraction Living in Nursing Homes

	Deaths Ages 65+ (age-adjusted)	
	women	men
BLS_UER	-0.001	-0.001
	(0.002)	(0.001)
UER X % Over 65 Living in Group Quarters	-0.056	-0.028
	(0.021)	(0.028)
	Deaths ages 0-45 (age-adjusted)	
	women	men
BLS_UER	0.004	-0.005
	(0.006)	(0.006)
UER X % Over 65 Living in Group Quarters	-0.113	-0.064
	(0.076)	(0.152)

Notes: Parameters are estimated Mortality semi-elasticity with respect to the state-year unemployment rate. Controls include State and Year Fixed effects, State-specific trends, demographic and education controls. Standard errors clustered at the state level. Estimates weighed by population. N = 1479.

Table 9

Cyclicality of Employment by Occupation: Skilled Nursing Facilities

	Total		Physicians		Nurses		Certified Aides		Other	
UE Rate	0.0316*	0.0375**	0.0138	0.0129	0.0214	0.0256*	0.0270**	0.0337**	0.0277**	0.0358**
	(0.0182)	(0.0182)	(0.0255)	(0.0241)	(0.0140)	(0.0132)	(0.0125)	(0.0135)	(0.0133)	(0.0143)
N=	185779	185779	56477	56477	185482	185482	182577	182577	184808	184808
Facility FE	X		X		X		X		X	
State FE		X		X		X		X		X

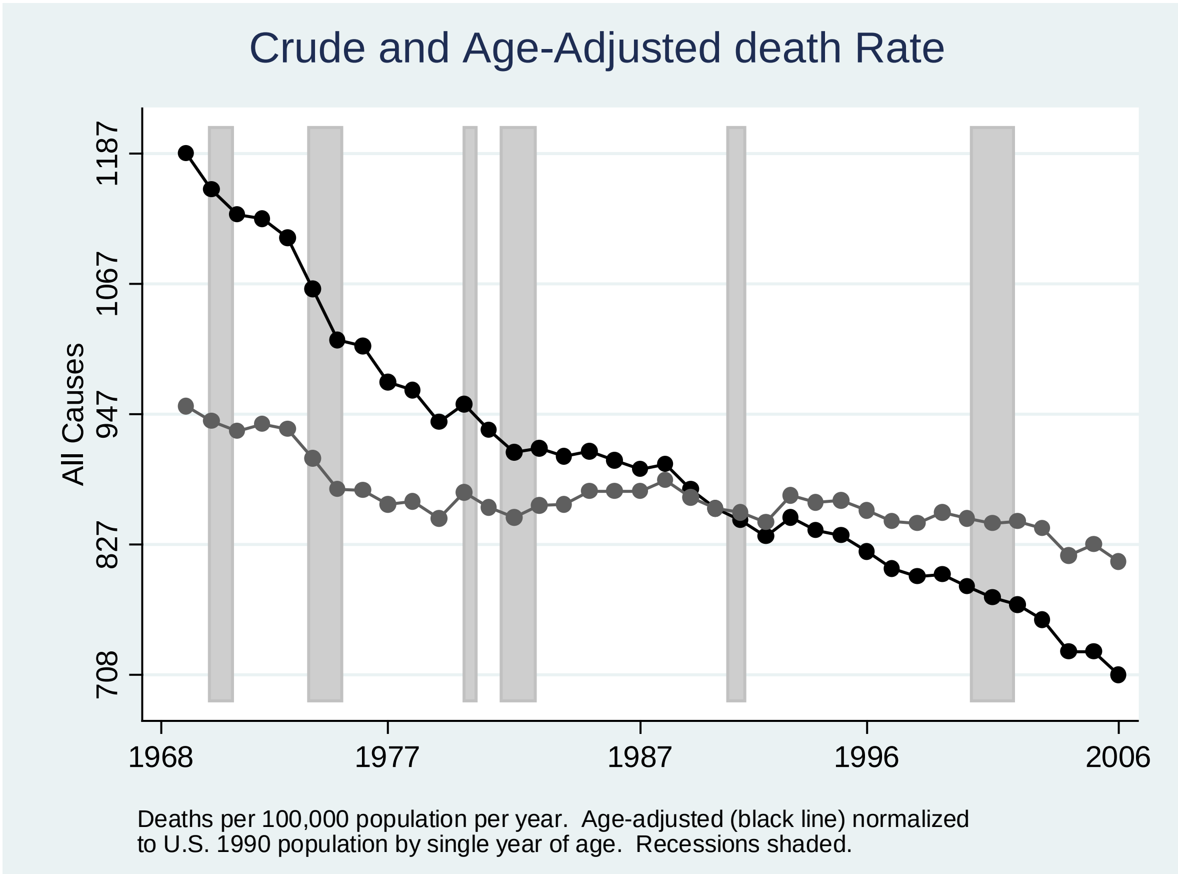
Linear regressions, dependent variable is log(employment count). Unit of observation is the facility-year, from OSCAR files, 1990 through 2006. All specifications include year and state fixed effects and state-specific trends. All models weighted by # of beds.

Table 10
Effects of the Unemployment Rate of Fraction of State Population Employed in Health Occupations

Occupation	Mean Fraction	Coefficient on Unemployment RateX100
MD	0.0032	-0.015
		(0.005)
Nurse	0.012	-0.009
		(0.011)
LPN	0.0022	-0.007
		(0.003)
RN	0.0094	-0.002
		(0.010)
Aides	0.0113	0.015
		(0.009)
Nursing Aides	0.0093	0.018
		(0.007)
Health Aides	0.002	-0.003
		(0.004)

Notes: Data from Monthly CPS, 1983-2002. The "Nurse" category is subdivided into LPN and RN; the "Aids" category is subdivided into "Nursing Aids" and "Health Aids".

Figure 1



Crude and Age-Adjusted death Rate
All Causes
1187
1067
947
827
708
1968
1977
1987
1996
2006
Deaths per 100,000 population per year. Age-adjusted (black line) normalized
to U.S. 1990 population by single year of age. Recessions shaded.

Figure 2

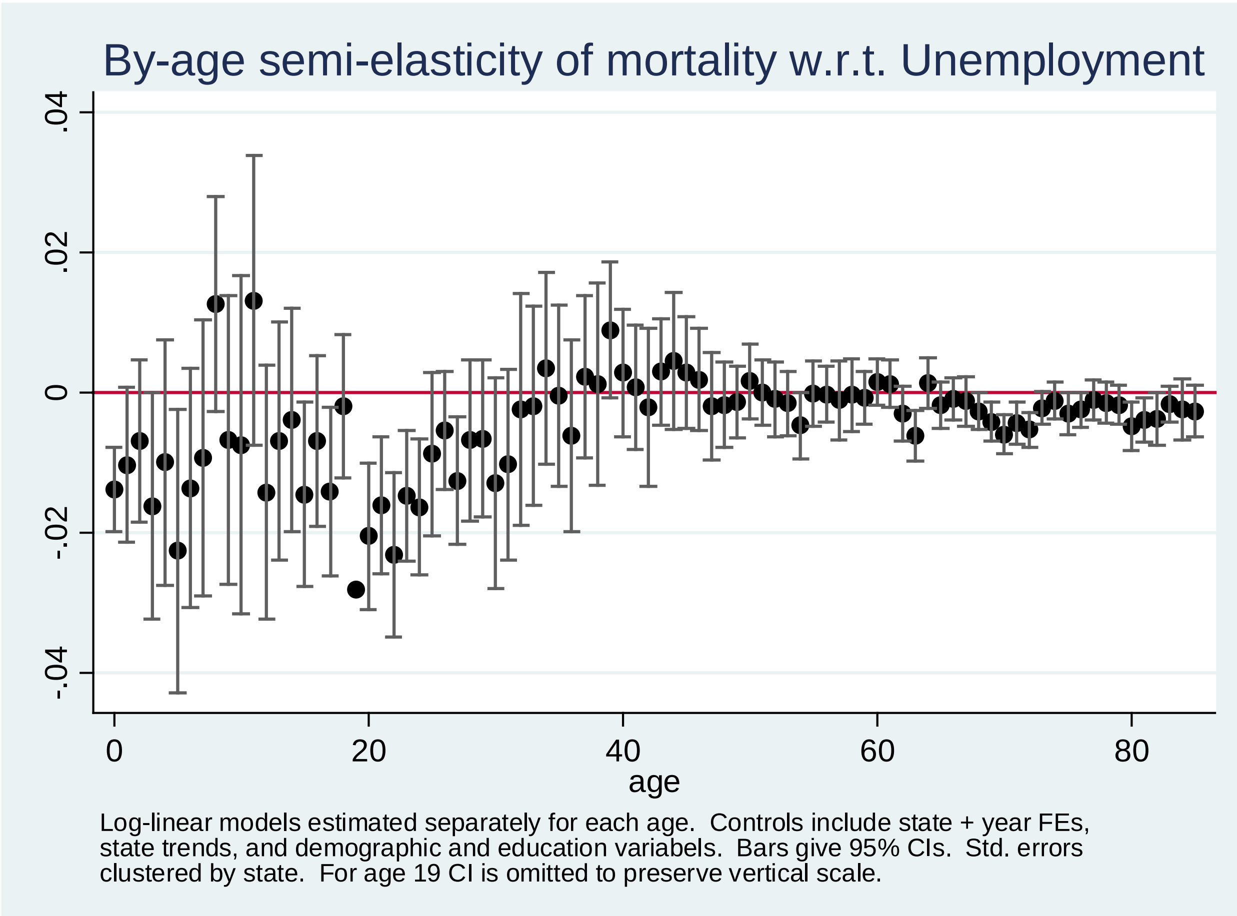
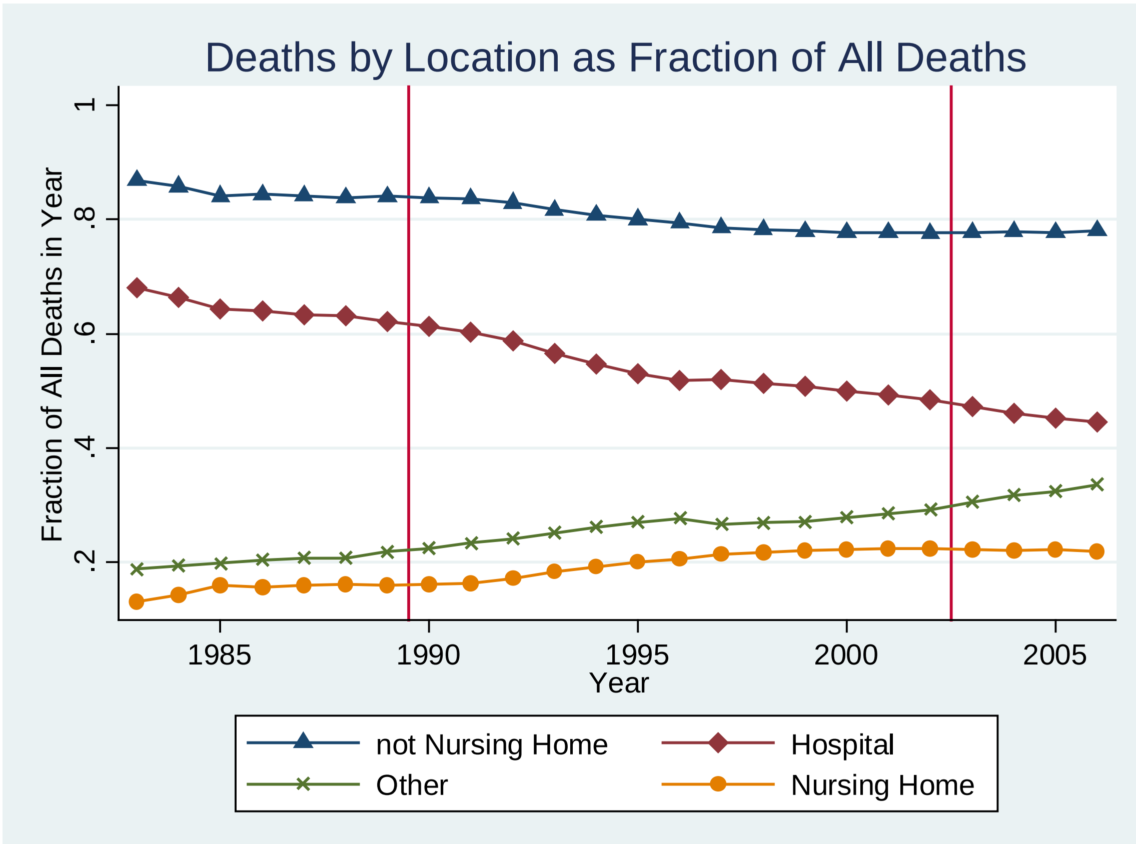

By-age semi-elasticity of mortality w.r.t. Unemployment
.04
.02
0
-.02
-.04
0
20
40
60
80
age
Log-linear models estimated separately for each age. Controls include state + year FEs,
state trends, and demographic and education variabels. Bars give 95% CIs. Std. errors
clustered by state. For age 19 CI is omitted to preserve vertical scale.

Figure 3



Note: Vertical lines at dates of change in death certificate reporting of location.