A Dynamic Model of Elderly Living Arrangements in Taiwan

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1. Introduction

An elderly person's living arrangement choice, much like his or her retirement decision, is a dynamic decision which should be responsive to changes in individual circumstances and to changing expectations about the future. However, evidence of strong persistence in living arrangements in countries such as the U.S. and Japan raises questions about whether living arrangements, in fact, can adjust quickly to changing conditions (Borsch-Supan, 1990; Brown and Liang, 2000). Policy concern over elderly living arrangements arises from the large effect of living arrangement choices on elderly care and welfare, especially for those suffering from physical limitations or health problems.[1] This concern may be especially cogent in rapidly developing countries, such as those in East Asia, where the percentage of elderly living independently (alone or with their spouse) has increased substantially but where public support services for the elderly remain relatively underdeveloped.

There are numerous possible causes for observed persistence in living arrangements. First, there may be state-dependence--being in a particular living arrangement is self-reinforcing. The elderly may resist changes in living style (habit formation), there may be significant costs (financial, time, emotional) to moving for parents or children, or difficulty in making other changes in work or family life necessitated by a new living arrangement. The degree to which different living arrangements are self-reinforcing may vary. Second, unobservable factors related to the health, preferences, wealth, work status, location, etc. of parents and children, or that affect the closeness of family relationships (e.g., past personal disputes), may persist over time. As in other contexts, the existence of such unobserved heterogeneity, if not controlled for, may result in biased estimates of state-dependency and the effects of other determinants. Finally, another aspect of persistence is the extent to which living arrangements respond directly to *changes* in specific circumstances such as health and marital status.

To analyze these dynamic aspects of the living arrangement decision requires panel data. Unfortunately, in contrast to work on the retirement decision, almost all empirical research to date on living arrangements has used cross-sectional data. Such work cannot distinguish between the effects on living arrangement choice of cohort differences and changes that occur with aging. For example, if health strongly affects living arrangements in a cross-sectional comparison, one cannot infer that living arrangements respond adequately to deteriorating health of individuals over time (a within rather than between effect). More generally, the rapid shift to independent living observed in many Asian countries may be attributable both to changes in cohort characteristics and to within-cohort changes over time combined with changing cohort composition of the elderly population.

Two previous studies using U.S. panel data are handicapped by data limitations. Borsch-Supan (1990) examines living arrangement transition probabilities using the PSID, but the data reveal few transitions, provide no information on elderly health and care, and have incomplete information about elderly individuals who are not household heads (e.g., number of children). Borsch-Supan et al. (1992) estimate a multiperiod model using survey data on the elderly in the Boston metropolitan area collected by the Hebrew Rehabilitation Center for the Aged (HRCA). While the data is rich, the sample used for estimation is small (314 elderly) and has "peculiar features" (Venti, 1992). Neither paper explicitly introduce dynamics into a model of living arrangement choice.

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[1] In Taiwan, elderly with functional limitations are much more likely to receive assistance and to be satisfied with their care if they live with children rather than live independently (Chen, 1994; Lin et al., 2000).

In this paper we exploit panel data from the Survey of Health and Living Status of the Elderly in Taiwan (SHLSET) collected by the Taiwan Provincial Institute of Family Planning in collaboration with the University of Michigan's Population Studies Center. The nationally representative survey of elderly above age 60 in 1989, was conducted in 1989, 1993, and 1996, and meets the considerable data demands for estimating a dynamic model—longitudinal data with information on health, income, wealth, and family demographics for a large group of elderly. The number of individuals surveyed in the three waves were 4049, 3155, and 2669.[2] Unlike the U.S. and Japan, there are considerable changes over time in living arrangements, with 32 percent of individuals surveyed in 1996 experiencing at least one change since 1989.[3] This provides sufficient variation for studying dynamics.

We utilize the Taiwanese panel data to estimate, we believe for the first time, a fully dynamic empirical model of elderly living arrangement choice that tests simultaneously for state-dependence, unobserved heterogeneity, and the responsiveness of living arrangements to changes in health and marital status. The model is estimated using the GHK simulation estimator to allow for error correlations across different living arrangement choices and across time periods. We also exploit the panel data to decompose living arrangements and elderly characteristics into between-cohort changes and within-cohort changes that accompany aging. The estimation results are used to conduct simulations that quantify the effects of covariates on living arrangement transitions.

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[2] The response rate in the first year was 92 percent, almost entirely due to an inability to locate individuals selected from household registers (TPIFP, 1989). Of the attrition from 1989 to 1996, 1016 (75 percent) were confirmed deaths, with the rest due to unconfirmed death, inability to locate, or refusal or inability to participate.

[3] Of those that survive from one period to the next, 20 percent change category from 1989 to 1993, 22 percent from 1993 to 1996 (Table 2). Elderly living with others are particularly likely to change their living arrangement (less than half stay). Regarding attrition due to mortality, although differences are not large, independent livers are the most likely to survive, followed by those living with children, those living with others, and those who are institutionalized, who are twice as likely to die than independent livers.

The dynamic analysis yields a rich and in some ways unexpected story of living arrangement trends. First, despite the rapid shift to independent living in Taiwan among the elderly in aggregate, we find relatively weak cohort differences for recent elderly cohorts, especially for men. However, we do find a strong within cohort trend toward independent living as the elderly age, even though aging is accompanied by failing health, the loss of one's spouse, and the marriage of children—all of which predict less independent living. This surprising trend results from an initial distribution of living arrangements in which most elderly live with children and state-dependence in observed living arrangements over time.

The dynamic features of the estimated model turn out to be important. Estimated state-dependence is substantial; the elderly living independently and with children are 50 percent less likely to switch than to maintain their current living arrangement. Unobserved heterogeneity measured by the variance of random effects is significant for men but not for women. Although independent livers tend to be educated, healthy, and wealthy, living arrangements do not adjust significantly to changes in health. The loss of a spouse, however, has a large effect on living arrangement choice. By comparing estimates from the dynamic model with those that do not allow for state dependence or unobserved heterogeneity, we show that estimates from specifications frequently used in previous literature are likely to systematically overestimate the importance of other covariates for living arrangement choice.

Finally, based on simulation results, we conclude that the government's planned universal pension is unlikely to have a major effect on elderly living arrangements.

The rest of the paper is organized as follows. In section two, we describe trends in elderly living arrangements in Taiwan, conducting simple age and cohort decomposition exercises. In section three, we present an estimable dynamic model of elderly living arrangements and discuss various empirical issues. Section four describes the variables used and predictions for how different factors are expected to affect living arrangement choices. Section five presents the main estimation results. Interpretation of the causes of aggregate trends toward independent living as well as between-cohort differences and within-cohort changes in living arrangements is offered in section six. Results of a policy simulation that introduces universal pensions are presented in section seven. Section 8 concludes.

2. The Shift to Independent Living in Taiwan

Many countries in Asia are now challenged by the rapid aging of their populations. In Taiwan, those aged 65 and older accounted for 8.1 percent of the population in 1997, compared to 5.5 percent a decade earlier and 3.0 percent in 1970 (TSDB, 1998). The elderly share of the population is expected to rise to 9.9 percent by 2010 and to 21 percent by 2035 (Hu, Chen, and Chen, 1999). Throughout Asia, population aging and rapid socio-economic change have been accompanied by substantial increases in the number of elderly living independently (alone or with their spouse) and declines in the percentage of elderly cohabiting with children.[4] The changes are especially pronounced in Taiwan, where the percentage of those 65 and older living independently grew from 9 percent in 1976 to 38 percent in 1996 (Figure 1). The Taiwan case thus provides a unique opportunity to study the shift in living arrangements both in its early stages and in a context of rapid change, elements lacking from studies of the Western experience.[5] Insights from Taiwan can inform our understanding of similar processes occurring or likely to occur elsewhere in Asia.

[4] The percent of elderly living with children fell from 77 to 60 percent from 1970 to 1989 in Japan, 71 to 56 percent from 1970 to 1984 in South Korea, 74 to 70 percent from 1984 to 1996 in the Philippines, 88 to 85 percent from 1988 to 1995 in Singapore, and 76 to 71 percent from 1986 to 1995 in Thailand (Hermalin, 1999).
[5] An exception is Costa (1997).

In contrast to the U.S., a majority of the elderly in Taiwan view living with adult children, in particular eldest married sons, as the ideal living arrangement. This heightens concern that independent living is dictated not by preferences but by constraints, leaving increasing numbers of the elderly isolated and without adequate care. Are Taiwan's independent elderly poor and neglected, or wealthy and progressive?

In Table 1, we present a breakdown of living arrangement choice from the SHLSET data. We know from Figure 1 that in aggregate, independent living has increased steadily over time. What is striking in Table 1 is that this trend appears to be true not just across but within cohorts. One might expect independent living to decrease with age because of greater need for care due to deteriorating health—the pattern observed in the U.S. (Borsch-Supan et al., 1992). But following the same group of elderly over time in the SHLSET sample, independent living increases from 22.3 percent in 1989 to 26.9 percent in 1993, remaining the same in 1996. This trend is not due to selection effects from differential mortality. Although independent elderly do have slightly lower mortality than other groups (Table 2), restricting the sample to those surviving to 1996 does not appreciably alter the magnitude of the increase in independent living (Table 1). Nor is the trend due to the decline in unmarried children (most of whom live at home); the trend is the same when looking at elderly with no unmarried children in 1989.

We decompose the trend to independent living into age and cohort effects by regressing dummy variables for living arrangement choices on a set of age and cohort dummies. In the left hand panels in Figure 2, we plot the propensity of different living arrangements for different age groups based on three specifications: 1) age group dummies only (*age*); 2) age group and year dummies (*year*); and 3) age group and cohort dummies (*cohort*), which also is plotted separately by gender (*men, women*). Consistent with the discussion above, when controlling for cohort effects, we find a strong positive relationship between age and independent living (and with living with others), and a declining propensity to live with children as one ages. When aged 80 and older, the probability of living independently is twelve percent higher than when aged 60-65 while the probability of living with children is 28 percent lower.

The cohort effects, plotted in the right-hand side panels of Figure 2, reveal surprisingly little evidence of large cohort differences over the past twenty years. For men, there is almost no change, while for women the most recent cohort is about 10 percent more likely to live independently than the three cohorts that preceded it (among which there is little difference). However, living arrangements are significantly different for the older cohorts. Compared to the 60-64 year olds in 1989 (born during 1925-29), those aged 85 and older in 1989 (born before 1905) are 15 percent less likely to live alone and 25 percent more likely to live with children. This effect is stronger for women than men. Those aged 80 and over comprise only 6 percent of the sample, however, so older cohort effects are estimated with less precision.

Although age effects appear to be more pronounced than cohort effects, the overall direction of both effects is consistent with secular changes in the propensity to live independently, or time effects. Whether old or young, over time everyone is more likely to live independently and not with children. This same conclusion has been made about the increase in Taiwan's private savings rate (Deaton and Paxson, 1998). If we disallow cohort effects in favor of time effects, we find a pattern of age effects very similar to those based on regressions with age dummy variables only--independent living declines and living with children increases after age 75 (Figure 1).

Transition probabilities presented in Table 2 show that while there is greater likelihood of moving in with children if one is living independently than vice versa (17 versus 10 percent), because the absolute numbers of independent livers are much smaller, the total number of individuals moving from living with children to independent living is twice as great as the opposite flow. This suggests that the large numbers of elderly living with children in the initial distribution of living arrangements may help explain the within-cohort trend to independent living.

3. An Estimable Dynamic Model of Living Arrangements

We define three living arrangement choices—living independently (alone or with spouse only), living with others, and living with children.[6] We have data for three years--1989, 1993, and 1996, which we denote as periods 0, 1, and 2.

We can write the underlying model for living arrangements in periods 1 and 2 as follows:

$$L_{it}^{k*} = \alpha^k L_{it-1} + \beta^k X_{it} + \eta^k X1_{it-1} + \tilde{\gamma}_i^k + \tilde{e}_{it}^k,\ \ k=d,o,c \tag{1}$$

$$\gamma_i^k = \tilde{\gamma}_i^k - \tilde{\gamma}_i^c,\ \ e_{it}^k = \tilde{e}_{it}^k - \tilde{e}_{it}^c,\ \ k=d,o \tag{2}$$

The initial condition (period 0) is the following:

$$L_{i0}^{k*} = \beta_0^k X_{i0} + \tilde{v}_{i0}^k,\ \ k=d,o,c \tag{3}$$

$$v_i^k = \tilde{v}_i^k - \tilde{v}_i^c,\ \ k=d,o \tag{4}$$

[6] We exclude individuals who lived in nursing homes or other institutions in any year, which account for less that 2 percent of the sample.

L_{it}^{k*} is the utility of living arrangement k for person i in period t, and L_{it}^{k} is an indicator variable for whether choice k is observed. Individuals choose the living arrangement with the highest utility. The k can take on three values: living independently (d), living with others (o), or living with children (c). Without loss of generality, we set the coefficients for the reference choice, living with children, equal to zero and interpret the coefficients for the other choices as the effects of covariates in comparison to their effects on the reference choice. Thus, equations for the utility difference with respect to the reference choice are the same as (1) and (3) except that the error components ($\tilde{\gamma}_i^k, \tilde{e}_{it}^k$) are replaced by the error differences (γ_i^k, e_{it}^k) defined in (2) and (4). L_{it} is a 2-element vector of the previous period living arrangement (dummies for living independently and living with others). The independent variables include both current and lagged variables, represented by X_{it} and $X1_{it-1}$, where X1 is a subset of X.

The model is dynamic in three respects. First, we introduce state dependence by including lagged living arrangement dummy variables (L_{it-1}). Second, we allow for unobserved heterogeneity by introducing random effects for each living arrangement choice (γ_i^k). Third, we include lagged independent variables for health and marital status, factors which change greatly over time and are likely to have large effects on living arrangement choice. This enables us to study the speed of adjustment and the importance of changes versus levels for living arrangement outcomes. To our knowledge, this is the first paper to fully incorporate dynamics into an empirical model of living arrangement choice. Borsch-Supan et al. (1992) allow for serial correlation (persistent shocks) but not state-dependence or lagged independent variables.

Our main interest is in the structural parameters (α^k , β^k, and η^k) generating the stochastic process in periods 1 and 2. Inclusion of lagged dependent variables leads to the problem of endogenous initial conditions, in our case the 1989 living arrangement. Assuming that the initial period living arrangement is exogenous or that the process is in a state of equilibrium is problematic (Heckman, 1981). Following Heckman (1981) and Hsiao (1986, p.171), we specify a reduced-form probability function for initial living arrangement which includes exogenous regressors for the pre-sample period (1989 and before). The reduced form error terms are allowed to be freely correlated with errors in the structural equations in periods 2 and 3 (correlations among v^k_{i0} and u^k_{i1} and among v^k_{i0} and u^k_{i2}, where $u^k_{it} = \gamma_i^k + e^k_{it}$). This is quite general in that it fixes the initial conditions without imposing any assumptions about the correlation between the error terms in the second and third periods. Identification, however, relies partly on the distributional assumption of joint normality among the error terms. This approximate solution performs well in Monte Carlo simulations (Heckman, 1981).

We specify the system of equations as a multinomial probit. Unlike the multinomial logit which assumes that choices exhibit independence of irrelevant alternatives (IIA), we allow for correlations among the random effects (γ_i^d and γ_i^o). In principle, one could also model the correlation among the time-varying unobservables (e_{it}^d and e_{it}^o), but in a short panel, separate identification of both correlations is problematic (Borsch-Supan et al., 1992). Nonetheless, we also estimate the model with both correlations and conduct a likelihood ratio test to compare its explanatory power with our preferred specification.

Our preferred specification (which we call the *full model*) includes lagged dependent variables, correlated random effects, and initial period equations with errors freely correlated with the errors of the structural equations of interest for periods 1 and 2. For comparison, we also report estimates from two other specifications (Appendix Tables 1 and 2). The first specification (*IIA*) ignores unobserved heterogeneity (no random effects) and assumes that the errors for different choices are uncorrelated. Although lagged living arrangement dummies are included, the assumption of no random effects implies that the lagged living arrangement is exogenous, so we need not model the initial period living arrangement. The second alternative specification (*no state dependence*) drops the lagged living arrangement variables but maintains correlated random effects, similar to the multinomial probit models estimated by Borsch-Supan et. al (1992).

The difficulty in allowing for so many correlated errors (across both choices and time periods) is that the likelihood function must integrate across many dimensions, which is computationally intractable. We assume that the correlated errors have a joint normal distribution, and estimate the model by employing the GHK simulation-based estimator (Geweke, 1989; Hajivassiliou, 1993; Hajivassiliou and McFadden, 1998; and Keane, 1994).

In addition to omitting the elderly who ever lived in an institution (less than two percent of the sample), to ensure that all living arrangement choices are feasible we exclude those who report having no adult children in any year. Our sample includes all remaining individuals with complete data in both 1989 (necessary to fix the initial condition) and 1993, as well as those with complete data in 1996. This yields a sample of 2623 in 1993 and 2101 in 1996, or 83 and 79 percent of the total samples. Because behavior is likely to differ significantly for men and women, all specifications are estimated separately by gender.

We note several additional empirical concerns. First, the discrete lag structure does not capture multiple transitions. However, changes in living arrangements are infrequent, so multiple changes are unlikely. In 1993, 93 percent of those reporting changing residence since 1989 report only one move, and some of those reporting more than one move rotate among children, so their living arrangement is not changing. In any case, residential change is not strongly associated with living arrangement changes. From 1989 to 1993, of 347 movers and 642 living arrangement changes, only 102 overlap. Second, the lag periods are not consistent (4 years and 3 years). Rather than making structural assumptions to adjust lag period differences, we interpret the lag coefficients as a weighted average of 3 and 4-year lag effects. Third, our model ignores selection effects due to differential mortality. Doing so would considerably complicate the estimation and Table 2 suggests that mortality rates do not vary substantially across living arrangements.

4. Determinants of Living Arrangement Choice

A number of factors influence living arrangement choices. In addition to the factors leading to persistence in living arrangements described in the introduction, explanations commonly offered for the shift to independent living include the following: 1) better health, which makes independent living feasible and more appealing; 2) fewer kinship resources due to demographic changes; 3) greater educational attainment, which increases the demand for privacy; 4) greater wealth, which makes separate living arrangements financially feasible; 5) government welfare and insurance programs, which substitute for private support for the elderly; 6) other secular factors that change over time, such as changing values, increasing opportunity cost of children's time, or changes in the housing market. Note that all of these factors (except 6) reflect cohort differences. In the U.S., the most conventional explanation for the shift to independent living is rising wealth and increasing demand for privacy. However, the effects of these factors on living arrangements may be different in Taiwan if preferences over living arrangements differ. In this section, we describe the variables used to test the validity of these different proposed explanations.

Persistence. State-dependence and unobserved heterogeneity may lead to persistence in living arrangements, preventing change even when other covariates favor a different living arrangement. The coefficients on the lagged dependent variables measure the extent of state-dependence in living arrangements. If we define α_{kl} as the effect of being in state k in one period on the utility of being in state l in the next period, then the coefficient on L_{it-1}^{d} in the equation for L_{it}^{d*} is equal to $((\alpha_{dd} + \alpha_{cc}) - (\alpha_{dc} + \alpha_{cd}))$ and the coefficient on L_{it-1}^{o} in the equation for L_{it}^{o*} is equal to $((\alpha_{oo} + \alpha_{cc}) - (\alpha_{oc} - \alpha_{co}))$, where c denotes living with children (see derivation in Appendix 1). The coefficients measure how likely one is to stay in one's starting state if the starting state is d or c (o or c) compared to the likelihood of switching from one state to the other. The coefficients thus jointly measure the state-dependent persistence of both the reference and choice living arrangements. Note that they do not identify whether it is harder to move from one state to the other or vice versa. The coefficients on the second lagged living arrangement variable in each equation, while having a less intuitive interpretation, also capture the relative persistence of a combination of living arrangement transitions.

As noted earlier, observed persistence can also be due to unobserved heterogeneity, which in our specification is measured by the estimated variances and covariances of the random effects. Finally, we can directly measure the responsiveness of living arrangements to changes in specific circumstances by including lagged independent variables, which we do for health and marital status.

Health. We measure health by the mean score for self-reported limitations in twelve activities of daily life (ADLs).[7] These functional health measures are more objective than self-reported health status and so are probably less endogenous to reporting bias. Nonetheless, they remain subject to measurement error because they do not capture all of the dimensions of health that are relevant for living arrangement choice. If such measurement error is large, our estimates may underestimate the importance of health and exaggerate the effect of economic variables (Bound, 1991).[8] However, in contrast to many economic outcomes such as labor force participation, living arrangements have been found to be affected more strongly by ADL measures than self-reported health status, suggesting that they better capture relevant dimensions of health for this particular outcome. This also accords well with our intuition that the physical presence of caregivers is most important for those with functional limitations. Other health problems may require more specialized medical services.

In our estimating equations we include a lagged health variable to examine health dynamics. If living arrangements respond quickly to changes in health, then conditioning on lagged health, current health should have a large effect on living arrangements. Lagged effects will enter significantly if adjustment is slow or if lagged health contains information about health persistence that is independent of current health.

Kinship Resources. We measure kinship resources with eight variables: marital status, marital status interacted with three spousal characteristics (age, education, work status),

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[7] ADL measures are scores from 0-3 for ability to do the following: crouch, reach above head, grasp, lift 25-pound object, walk 200-300 meters, climb 2-3 flights of stairs, take bus or train, do heavy work around house, bath, make a telephone call, buy personal use items. The score definitions are no difficulty (0), some difficulty (1), significant difficulty (2), and require assistance (3). Results using a weighting scheme based on factor analysis produced nearly identical results.

[8] One solution to the endogeneity and measurement error problems of health measures is to instrument self-reported health status with more objective health measures; but this will underestimate the effect of economic variables and can even exacerbate bias from using self-reported measures alone.

whether the respondent has any adult sons, the number of married adult sons, the number of married adult daughters, and the number of unmarried adult sons or daughters. Spouses provide financial, physical, and emotional support, and so their presence can have a large effect on the willingness to live alone. To look at dynamics, we include a lagged marital status variable.

Education. Years of education is correlated with wealth, and also has a strong independent positive effect on preferences for independent living (Kan and Park, 2000).

Wealth. Different dimensions of wealth are measured by four variables: work status, whether respondents or their spouses have pensions, whether they own property (housing or real estate), and whether they have divided property among children in the past.[9] In Taiwan, it is common for the elderly to divide up much of their wealth (especially agricultural land) before death—a kind of pre-death bequest. Work status and pensions are direct income sources and contribute to wealth.

Although wealth is positively associated with independent living in the West, in the East the relationship is less obvious if many elderly prefer to live with their children. If the elderly prefer to live independently, wealth likely plays a facilitating role. However, if the elderly prefer living with children, a power bargaining model would predict that the wealthy get their way (children cohabit to increase their inheritance). An altruistic model, on the other hand, suggests that with less need, the elderly get less assistance, via cohabitation, from their children (Lee, Parish, and Willis, 1994).

[9] We tried other wealth-related variables such as whether the respondent reports financial savings or other large assets (such as business fixed assets), but these were insignificant in all regressions and were excluded for parsimony. The survey also asks about income, but the income measure is not asked in a consistent way in the different survey years. Also, because many respondents no longer work, current income is a very noisy measure of permanent income or wealth.

We recognize the potential endogeneity problem with the property variables if living arrangement decisions are made simultaneously with the decision whether or not to divide property to kids, for example if property is divided as part of an agreement in which children agree to live with their parents. If this is the case, property would be associated with independent living and property division with living with children, and the joint result would support an altruism or mutual exchange model of living arrangements. If both property and property division increase the likelihood of independent living, wealth has a conventional effect, either by facilitating or creating a preference for independent living or reducing the pressure for children to care for parents through cohabitation.

Public Assistance. Pensions in Taiwan are provided by the government, so the pension variable can be considered an indicator of public assistance. The variable has particular policy relevance in light of the government's announced plan to introduce a universal pension program for all citizens, which follows the establishment of a national health insurance program in 1995.

Government pensions are linked closely to occupational categories. While this may introduce some bias due to the endogeneity of occupational choice, we have no strong priors that lead us to expect that unobserved factors affecting occupational choice will be systematically correlated with living arrangement choices after controlling for education, wealth, health, and kinship resources. The pension variable could also pick up the effect of other government benefits tied to occupational choice, such as health insurance.

Time Effects. A year dummy for 1996 captures time-related changes, such as in popular attitudes toward intergenerational relations, housing prices, labor market conditions, and unobserved improvements in information, support services, etc., that could influence living arrangement choices.

5. Results

To provide an initial sense of what types of elderly choose different living arrangements, we compare sample means of elderly characteristics for each living arrangement group (Table 5). Men and women living independently on average have 6.2 and 2.5 years of schooling, compared to 5.2 and 1.6 years for those living with children and 5.5 and 1.7 years for those living with others. For both men and women, those living independently are healthier, more likely to be working, have spouses that are more educated and more likely to work, and have more property and pensions. They also have as many married children (for men) or more (for women) than those living with their children. Those living with others have the poorest health, yet are more likely to work (especially women), and have fewer children.

In Table 6, we report coefficient estimates for the full model. We also estimated a more general model that also allows both the random effects and the time-specific errors in the living arrangement choice equations to be correlated. Estimated standard errors in the general model are very high, suggesting that the model is not well-identified. Likelihood ratio tests (not reported) find that the "general" models (for men and women) do not explain additional variation in the data when compared to the full models. However, likelihood ratio tests do find that the full models explain significantly more variation than models that do not allow for any error correlation.

Persistence. The large and statistically significant coefficient for the lagged independent living dummy variable in the independent living equation for both men and women reflects a high degree of state dependence in the decision whether to live independently or with children (Table 6). For these two states, the difference in the marginal probability of staying versus switching is 51 percent for men and 50 percent for women

(Table 8). The comparable marginal probabilities for the decision to live with others versus with children is 16 and 18 percent for men and women, suggesting that living with others is a much less persistent state--a result consistent with the transition probabilities in Table 2.

Failure to allow for state dependence by including lagged living arrangement dummy variables leads to substantial exaggeration of the effects of covariates on living arrangement choices (see Appendix Table 2). For example, the coefficients on health (mean ADL score) increases in magnitude from –0.106 to –0.163 for men and from -0.132 to –0.312 for women; the education coefficient increases from 2.7 to 4.5 for men and from 1.2 to 2.7 for women. Thus, earlier studies which do not explicitly model dynamics may substantially overestimate the responsiveness of living arrangements to changing circumstances or policies.

The importance of unobserved heterogeneity as measured by the estimated variances of the random effects are large and statistically significant for men, but small and statistically insignificant for women (Table 7). The correlation between the random effects (for living independently and living with others) are positive, small, and insignificant for men, and much larger but still insignificant for women. Overall, the results suggest that unobserved heterogeneity is a greater concern for men than women, and that correlations among unobservables affecting different living arrangement choices are not significant.

There is also evidence that the error terms in the initial period equations are correlated with those in the living arrangement equations for future periods. Specifically, the correlations between living independently in 1989 and living independently in future years are large and positive (from 0.31 to 0.49) except for the correlation between 1989 and 1996 for women. The error correlations over time for living with others, in contrast, are negative and smaller in magnitude (-0.10 to –0.27), reflecting the more transitory nature of living with others. The inter-choice correlations across different periods are generally smaller and not precisely estimated, but tend to be larger for women. Overall, the results justify our effort to separately specify the initial living arrangement equations, and provide additional evidence of the importance of unobserved heterogeneity. Ignoring these correlations would lead to exaggerated state dependence.

When random effects are omitted, as in the multinomial probit IIA model estimates (Appendix Table 1), the measured persistence of living independently and living with children (coefficient of L_{it-1}^{d} in independent living equation) increases in magnitude because the lag terms pick up the effects of both state dependence and unobserved heterogeneity. The coefficient on the lagged independent living dummy in the independent living equation increases from 1.48 to 1.74 for men and from 1.66 to 1.82 for women when the random effects are dropped.

Health. Current functional limitations increase the likelihood of living with children rather than living independently but have little or no effect on the decision to live with others versus with children. While the coefficients on the mean ADL score is more negative for women (-0.132 versus -0.106, Table 6), the marginal effects are larger for men when evaluated at sample means (-0.030 versus –0.019, Table 8). This contrast occurs because women report significantly more functional limitations than men (Table 3). Although the mean ADL scores nearly double from 1989 to 1996 (0.345 to 0.600 for women and 0.141 to 0.312 for men, Table 3), given the small marginal effects, the health changes reduce the average predicted probability of living independently versus with children by less than one percent. As expected, lagged health also negatively affects the likelihood of living independently, but the coefficients are even smaller in magnitude and not statistically significant. We infer that the three-year period between survey rounds is sufficiently long for living arrangements to adjust to changes in health, although the adjustments are not large. Lagged health does not contain information about future health that affects living arrangement choice.

Kinship Resources. If an elderly man has an adult son, the likelihood that he lives independently versus with children falls by 18 percent (Table 8). Surprisingly, however, there is no effect for women. However, only about five percent of the sample lack adult sons. For both men and women, having an unmarried child reduces the probability of independent living by 5 percent. The number of married sons and daughters has little effect on changes in living arrangements. Thus, changes in the number of children are unlikely to explain changes in living arrangements, a finding consistent with Chen (1992).

The most important kinship resource affecting an individual's living arrangement is whether he or she has a spouse. While the coefficient for current marital status is not statistically significant, the interaction terms between marital status and spouse's education (as well as spouse's work status for men) have large and statistically significant coefficients (Table 6). For those who are married, an additional year of education of the spouse increases the probability of independent living by about 0.4 for both men and women (Table 8). The marginals for lagged marital status suggest that men and women who lose a spouse are four percent less likely to live independently than those who did not have a spouse in the previous period (approaching statistical significance only for women). The importance of marital status confirms observed trends in the data. From 1989 to 1996, the increase in the share of elderly couples living independently is substantial (21.2 to 28.3 percent for male respondents and 21.4 to 33.7 percent for female) while the increase for single elderly is very small (for women) or negative (for men).

Education. Education has a very large effect on elderly living arrangements, especially for men. The marginal probabilities associated with an extra year of education evaluated at sample means are 0.75 for men and 0.18 for women (Table 8). Results for spouses also suggest greater importance of men's education; spouse's education is much less important than own education for men but much more important for women. The large effects may reflect the strong effect of education on preferred living arrangements (Kan and Park, 2000). Education also may be correlated positively with unobserved wealth.

Wealth. Owning property increases the likelihood that women live independently by 7 percent, but has no measurable effect for men. Dividing property reduces the probability of independent living for men by 3 percent, but has no effect for women (Table 8). This latter result weakly supports an altruism or mutual exchange model rather than a power bargaining model. Working increases the probability of independent living much more for women than for men (6 versus 2 percent), and increases the likelihood of living with others by 2 percent for both men and women. Overall, wealth effects seem significant for women but not for men.

Public Assistance. Having a pension increases a man's probability of independent living by 6 percent compared to living with children, but the effect is less than one percent for women (Table 8). Interpreted as a wealth effect, this result contrasts with the gendered results for other wealth variables. However, it could reflect the greater importance of pensions to ensuring financial independence and the greater pension entitlements of men versus women (since the pension variable includes both own pensions and those of the spouse).

Time. The coefficients on the time dummy variables (for 1996) are negative, and for men statistically significant. Thus, there is no evidence that secular changes are a main cause for the shift to independent living.

7. Why the Shift to Independent Living?

If both the size of age cohorts and life expectancy were constant, increases in the aggregate share of elderly living independently would have to be due to cohort differences. However, in reality, more recent cohorts are larger and life expectancy has increased steadily. From 1970 to 1997, life expectancy increased from 67 to 72 for men and 72 to 78 for women. Overall, the share of the elderly (over 65) older than 70 increased steadily from 53 percent in 1980 to 57 percent in 1989 to 63 percent in 1998. Aging of the elderly population and a positive relationship between aging and the likelihood of independent living, whether due to changes in living arrangements or differential mortality, could explain some of the aggregate trend toward independent living irrespective of cohort differences.

One way to better understand the relative importance of these different factors in explaining aggregate trends is to decompose aggregate measures of independent living by the elderly (over age 65) in 1989 and 1996 (Table 9), incorporating data on 65- and 66-year-olds in 1996 from a new 1996 SHLET survey of those aged 50-66. In each year, we divide the elderly into two age groups—younger (65-71) and older (72+). Over these seven years, the share of elderly men and women aged 65 and over living independently increased by 2.3 and 7.2 percent, respectively. Consistent with earlier results, recent cohort effects (comparing the younger age group in 1989 and in 1996) are not pronounced for men but are for women. For the older age group, cohort effects are significant for men and women, but especially for men. The share of the older group increases for both men and women and the average age of the older group increases slightly (by 0.1 years for men and 0.6 years for women). However, because of the slight difference in the independent living shares among younger and older groups for men and the strong cohort differences among women, it is clear that the aging of the elderly combined with within-cohort trends toward independent living cannot explain the aggregate trends. Nor does differential mortality contribute to increased independent living as the elderly age. In fact, for men, those living independently in 1989 are less likely to survive to 1996. Thus, despite the lack of strong recent cohort effects for men, it is still true that the aggregate trend toward independent living is mainly explained by cohort differences. At the same time, explaining within cohort trends remains important for understanding the welfare of the elderly.

Cohort Differences

We observed in section 2 that for independent living, cohort effects are not strong among men, except for much older cohorts, and are stronger but uneven for women (Figure 2). This is somewhat surprising given our strong priors that changes in education, health, and wealth should drive the shift to independent living, and the difficulty of finding factors that might lead to increasing independent living within cohorts.

To decompose cohort differences, following the earlier analysis of living arrangement trends, we distinguish between cohort and age effects by running regressions of the independent variables on age and cohort dummy variables. For variables that are time-invariant, we regress on cohort dummies only. The results, presented in Table 4, confirm significant cohort differences in education, pensions, and property division (more recent cohorts are more likely to divide property) for both men and women. The number of children of men falls in more recent cohorts, mainly because of a lower likelihood to ever marry (many men immigrated from mainland China following the Communist victory there).

However, in contrast to the conventional story of declining kin resources, the number of children of women increases in more recent cohorts, perhaps because the earliest cohort in the sample bore children before the rapid fertility decline in the 1960s. Several other cohort differences are more significant for women than men. Women in recent cohorts are more likely to have a spouse, likely a consequence of rising life expectancy of men. They also are more likely to own property, and have better health. While these trends are also true for men, they are much less pronounced. There is no evidence of cohort differences in labor force participation for either men or women. The gender differences in cohort effects helps explain why recent cohort differences in living arrangements are more pronounced for women than men. Factors such as marital status and property have relatively high marginal effects on independent living and are greater for women than for men.

Age Differences

Cohort differences cannot explain the increase in independent living in the SHLSET data, which follows the same group over time. As for cohort effects, variables influencing within-cohort changes in living arrangement should be those that change significantly with age and have large marginal effects.

A simple first step is to compare sample means for time-varying covariates in 1989 and 1996 to see how elderly characteristics change with age (there will be some bias in such a comparison due to selective mortality). Over the seven years, functional limitations double in frequency, a significant percentage of elderly lose their spouse (especially women), the majority of those working in 1989 retire, many elderly divide property among children, and about one fourth of unmarried children marry (Table 3).

Next, we examine the regression-based decomposition of cohort and age effects (Table 4). The results confirm the changes associated with aging found in the sample mean differences between 1989 and 1996. As people get older, their health deteriorates, they stop working, they divide property among kids, and children marry (Table 4). Women stop working sooner and their health deteriorates more rapidly.

To quantify which variables most affect within-cohort living arrangement changes, we conduct a simulation exercise. First, for each respondent we predict the probability of each living arrangement choice for 1993 conditional on the 1989 living arrangement and 1993 covariates. Then we predict the 1996 living arrangement probabilities by taking a weighted average of the 1996 probabilities over the three different 1993 living arrangement outcomes using the 1993 predicted living arrangement probabilities as weights. We run simulations in which we hold the value for each time-varying variable constant at its 1989 level in 1993 and 1996 (one at a time, and in combination), and then compare these to a baseline simulation. The reported predictions (Table 10) are the mean predicted probabilities.

Not surprisingly, with the exception of unmarried children, all time-varying factors predict less rather than more independent living over time. The decline in unmarried children explains only a 1.97 percent increase in predicted independent living for men and 1.53 percent for women, much less than the 6.10 percent and 5.04 increase in the empirical distributions over the sample period. For almost all variables, holding values constant at 1989 levels alters the predicted distribution of 1996 living arrangements by less than one or two percent. The major exception is marital status. The simulation, which assumes that those with spouses do not lose the spouses in future years (nor do spouse characteristics change), produces a result in which the predicted percentage of those living independently more than doubles for both men and women. The large effect of marital status is not surprising when one considers that forty percent of women and twenty percent of men lose their spouse from age 60-64 to 80-84 (Table 4).

Finally, if all time-varying factors are held constant at their 1989 levels, the predicted percentage of men and women living independently in 1996 increase substantially from their 1989 levels, to 50.8 and 45.0 percent, respectively. This illustrates the large negative cumulative effect of changes associated with aging on the likelihood of living independently.

We are left with the question of what can explain the within-cohort trend toward independent living. Although the simulations do not identify a smoking gun--a time-varying covariate that can explain the change, the model as a whole successfully predicts greater independent living for the sample over time, even though it under-predicts the magnitude of that shift.[10] This predictive power must be coming from the other covariates in the model, namely the lagged living arrangement variables.

We argue that the explanation for greater independent living over time lies in the persistence of living arrangement choices and the initial distribution of living arrangements. When children become adults and the family must decide on the living arrangement of the elderly, almost all elderly first live with their children even if their characteristics predict a strong predilection for independent living. Given strong state-dependence in living arrangement choices, for a given cohort the system moves only slowly to an equilibrium in which a sizable fraction of elderly live independently. This movement toward greater independent living occurs despite the fact that many time-varying factors are making independent living less attractive over time.

We make several observations in support of this interpretation. First, the fact that our estimated model predicts the trend toward independent living but that changes in the

[10] We believe under-prediction likely reflects heterogeneity in the effect of covariates that is not modeled.

covariates are not the source of this change strongly suggests that the accuracy of the predictions lie in the lag variable coefficients and the initial distribution of living arrangements. Second, transition probabilities reported in Table 2 show that the probability of living with children conditional on living independently is greater than the probability of living independently conditional on living with children, and yet there is a net movement of people from living with children to living independently because there are so many more people living with kids. If transition probabilities were stable, this suggests that the starting distribution has fewer independent livers than the "equilibrium" distribution. Third, in other work examining the relationship between stated preferences over living arrangements and actual living arrangements (Kan and Park, 2000), we find that the elderly who prefer living independently are much less likely to realize their preference than those who prefer living with children.

To see more clearly how trends in living arrangements depend on the initial distribution of living arrangements when there is state dependence, we conduct simulations in which we vary the initial distribution of living arrangements in our sample and see how these changes alter predicted distributions of future living arrangements. We use the predicted probabilities from the 1989 living arrangement choice equations to order individuals by their probability of living independently. Starting with the actual distribution of living arrangements (22.83 percent of men and 16.84 percent of women living independently), we increase the initial percentage living independently by reassigning the initial living arrangement of those not living independently, starting with those with the highest predicted likelihood of living independently. We also decrease the initial percentage living independently by reassigning alternative living arrangements to those living independently, starting with those with the lowest predicted likelihood of living independently. For the latter, the new living arrangement (with others or with children) is that with the higher predicted 1989 probability.

The results of this exercise are summarized in Figure 3, which plots the 1996 predicted percentage of men and women living independently against the simulated 1989 percentages living independently. The predicted percentage of those living independently is uniformly higher for men than women (the difference averaging about 5 percent). When the line is above the 45 degree line, the predicted 1996 percentage living independently is greater than the 1989 initial percentage. However, below the 45 degree line, the 1996 predicted percentage living independently is lower. The direction of change in living arrangements thus depends on the initial distribution of living arrangements. The points of intersection with the 45 degree line, at about 0.24 for men and 0.18 for women, are "stable" distributions which are threshold points for whether independent living is expected to increase or decrease over time. These stable points are both above the actual 1989 distributions, which explains why independent living increases as the elderly age despite changes in time-varying covariates that predict less independent living.

Although we are confident that the explanation for the within-cohort trend toward independent living lies in the very state dependence that is excluded in earlier empirical models of living arrangement choice, we have yet to provide much economic intuition. As children become adults, they start by cohabiting with their parents. Even after marriage, many families find it economical to live together. As both parents and children accumulate wealth they are more likely to make arrangements to live separately. In some cases, children may decide to migrate to urban areas, with parents choosing to stay if resources permit (Chen, 1992). Given the rapid income growth seen in Taiwan, wealth accumulation by children is likely to be more important than wealth accumulation by parents. Liu (1995) finds that child characteristics, especially wealth, matter much more than parental characteristics in predicting independent living. Lee, Parish and, Willis (1994) find son's wealth to be positively associated with separate living. The birth of grandchildren also may eventually create pressures on housing space, especially as they grow older (Chen, 1993). These factors, in addition to many other factors contributing to inertia, may explain persistence in living arrangements but an eventual shift to independent living arrangements. Further research testing such theories with appropriate data on children will be of great interest.

8. Policy Experiment: Universal Pensions

To examine the potential effect of universal pensions, a program which is being prepared by the Taiwanese government, we run a simulation in which all individuals in the sample have pensions. Conditional on the 1989 initial living arrangement, adding pensions increases the predicted percentage of those living independently by 1.78 percent for men and by 0.31 percent for women. This prediction assumes that universal pensions offer the same benefits as existing programs. The reason for the small change is twofold. First, most elderly in the sample already have pensions (72 percent of men and 58 percent of women in 1996). Second, the marginal effects, especially for women, are not very large. These results suggest that the effects of pensions on living arrangements will be limited.

9. Conclusion

In this paper, we exploit a remarkable panel dataset of the elderly in Taiwan from 1989 to 1996 to estimate, for the first time, a fully dynamic model of elderly living arrangement choice. We also decompose changes in living arrangements and elderly characteristics into cohort and age effects and conduct a set of simulation exercises to quantify the effect of different covariates on living arrangement transitions. Estimation of the dynamic model leads to new insights into the persistence of living arrangements and the responsiveness of living arrangements to changes in life circumstances facing the elderly. The results also provide a new understanding of the rapid trend toward independent living among Taiwan's elderly.

One important unexpected finding is that the shift to independent living occurs not just between cohorts but within cohorts as well. As the elderly age, they are more likely to live independently despite deteriorating health and the loss of spouses. This is because most elderly initially live with children, and even those with a strong predilection to independent living find it difficult to change living arrangements because of strong persistence in living arrangements. For men, unobserved heterogeneity explains some of this persistence, but most of it is attributable to state-dependence (for both men and women). While living arrangements appear to respond to the loss of a spouse, the responsiveness to changes in ADL limitations is not very large.

In cross-section, the more educated, wealthier, and healthier elderly are more likely to live independently, reducing concerns that the independent elderly tend to be poor and neglected. Nonetheless, the strong persistence and limited responsiveness of living arrangements to changing health status may carry significant welfare costs and justify policy concern. Although children who do not live with parents often substitute other types of support such as transfers and visits (Lee, Parish, and Willis, 1994; Chen, 1993), care for those in poor health is often insufficient when children do not co-reside with parents (Chen, 1994).

The aggregate shift toward independent living is mostly attributable to cohort differences. Secular changes (time effects) if anything are inhibiting the shift. Nonetheless, we find, interestingly, that for men recent cohort effects are not very pronounced, making it unclear how sustained will be the rapid shift to independent living. Recent cohort effects are larger for women, due to significant differences in the marriage rate, education, health, and property holdings.

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Appendix
Derivation of Coefficients of Lagged Living Arrangement Dummy Variables

The utility of each living arrangement can be expressed as a function of previous living arrangements:

$$L_{it}^{d*} = \alpha_{dd} L_{it-1}^{d} + \alpha_{od} L_{it-1}^{o} + \alpha_{cd} L_{it-1}^{c}$$

$$L_{it}^{o*} = \alpha_{do} L_{it-1}^{d} + \alpha_{oo} L_{it-1}^{o} + \alpha_{co} L_{it-1}^{c}$$

$$L_{it}^{c*} = \alpha_{dc} L_{it-1}^{d} + \alpha_{oc} L_{it-1}^{o} + \alpha_{cc} L_{it-1}^{c}$$

Given the identity that the three living arrangement dummy variables must sum to one, we can substitute for one of the dummy variables (constant term ignored):

$$L_{it}^{d*} = (\alpha_{dd} - \alpha_{cd}) L_{it-1}^{d} + (\alpha_{od} - \alpha_{cd}) L_{it-1}^{o}$$

$$L_{it}^{o*} = (\alpha_{do} - \alpha_{co}) L_{it-1}^{d} + (\alpha_{oo} - \alpha_{co}) L_{it-1}^{o}$$

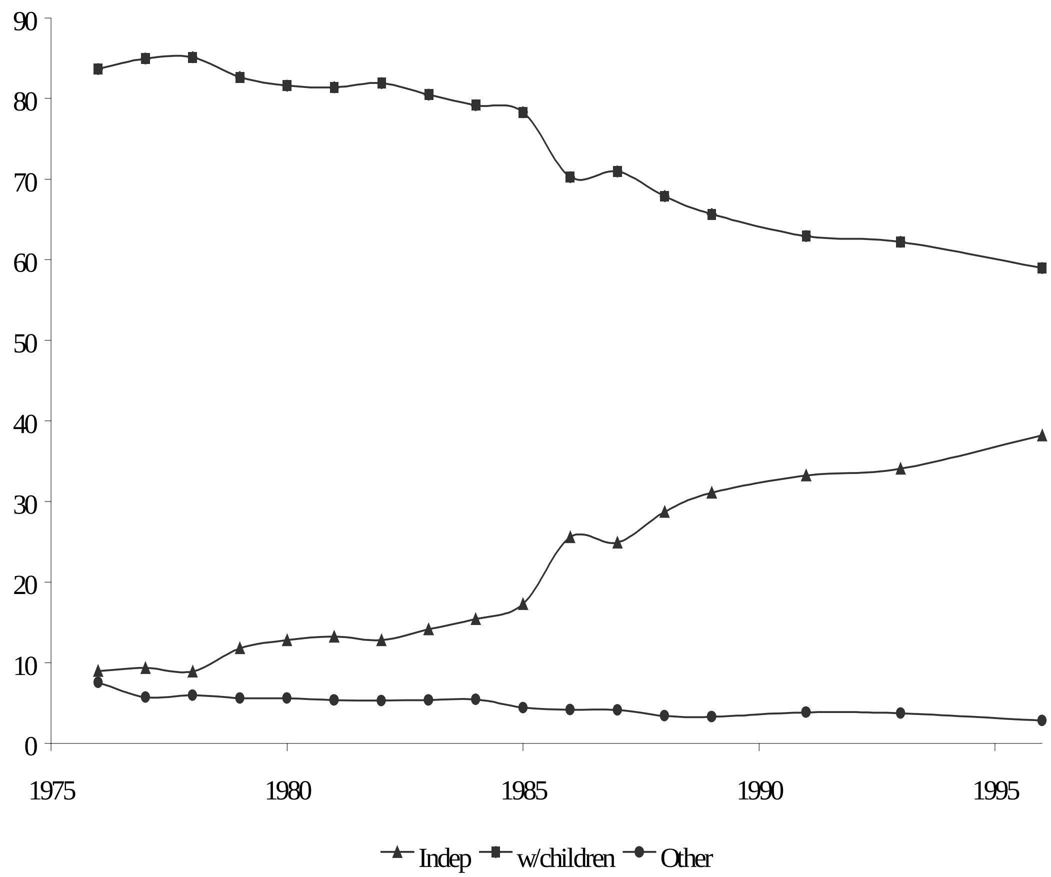
$$L_{it}^{c*} = (\alpha_{dc} - \alpha_{cc}) L_{it-1}^{d} + (\alpha_{oc} - \alpha_{cc}) L_{it-1}^{o}$$

The utility of living arrangement choices then can be expressed with respect to a reference choice (c):

$$L_{it}^{d*} - L_{it}^{c*} = ((\alpha_{dd} - \alpha_{cd}) - (\alpha_{dc} - \alpha_{cc})) L_{it-1}^{d} + ((\alpha_{od} - \alpha_{cd}) - (\alpha_{oc} - \alpha_{cc})) L_{it-1}^{o}$$

$$L_{it}^{o*} - L_{it}^{c*} = ((\alpha_{do} - \alpha_{co}) - (\alpha_{dc} - \alpha_{cc})) L_{it-1}^{d} + ((\alpha_{oo} - \alpha_{co}) - (\alpha_{oc} - \alpha_{cc})) L_{it-1}^{o}$$

Figure 1
Taiwan Trends in Living Arrangements of the Elderly Age 65 and Above
(in percent)



Sources: 1976-85 derived from Lo (1987). 1986-1996 data from the *Report on the Old Status Survey*, by Directorate-General of Budget, Accounting, and Statistics (DGBAS), Executive Yuan, based on surveys of over 1 million elderly. 1996 values interpolated based on new classification system.

Figure 2
Living Arrangement Age and Cohort Effects

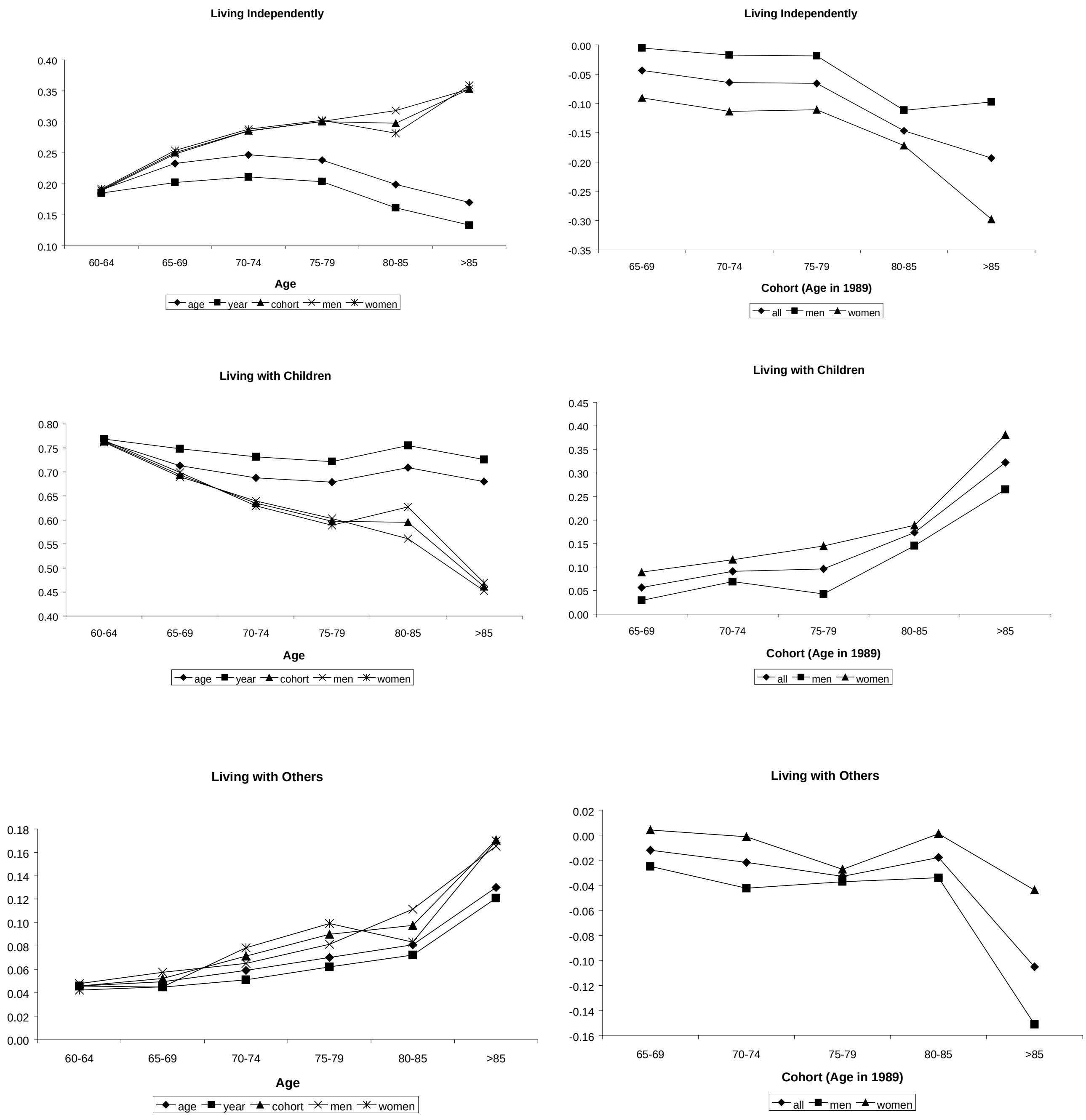


Figure 3
Simulated 1996 Independent Living Share Conditional on 1989 Independent Living Share

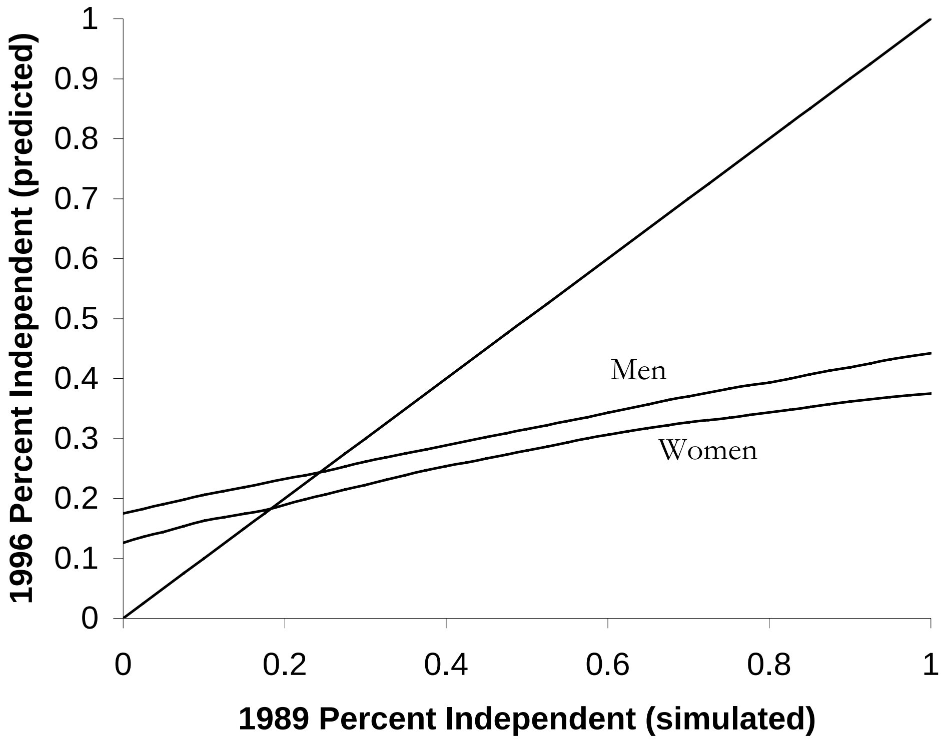


Table 1
Living Arrangements of the Elderly, 1989, 1993, and 1996
(Age 60 and Above in 1989)

	Full Sample			Balanced Panel Sample		
	1989	1993	1996	1989	1993	1996
Independent	22.3	26.9	26.9	21.8	26.8	26.7
With Children	71.0	65.9	63.4	72.4	66.8	64.0
With Others	5.8	5.9	8.2	5.2	5.5	8.0
Institution	1.0	1.3	1.5	0.6	1.0	1.4
N	4049	3155	2669	2525	2525	2525

Source: Survey of Health and Living Status of the Elderly, full sample means.

Table 2
Elderly Living Arrangement Transitions, 1989 to 1996

	Unit	Independent	W\children	W\others	Institution	Death
Independent	#	408	132	36	14	204
	%	51.39	16.62	4.53	1.76	25.69
With Children	#	258	1520	135	6	759
	%	9.63	56.76	5.04	0.22	28.34
With Others	#	50	38	47	8	65
	%	24.04	18.27	22.6	3.85	31.25
Institution	#	2	1	2	12	17
	%	5.88	2.94	5.88	35.29	50

Source: Survey of Health and Living Status of the Elderly.

Table 3
Sample Means for Men and Women, 1989 and 1996

		Men				Women			
		1989		1996		1989		1996	
Variable	Definition	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
		N=1358		N=1082		N=1265		N=1019	
Live independently	1=yes, 0=no	0.228		0.289		0.168		0.219	
Live with children	1=yes, 0=no	0.730		0.635		0.775		0.691	
Live with others	1=yes, 0=no	0.041		0.076		0.057		0.090	
ADL limitations	Mean score (0-3)	0.141	0.340	0.312	0.608	0.345	0.492	0.600	0.728
Married	1=yes, 0=no	0.822		0.704		0.542		0.385	
Age	years	67.2	5.8	73.7	5.5	67.9	6.1	74.3	5.7
Education	years	5.47	4.59	5.55	4.59	1.74	3.17	1.87	3.26
Work	1=yes, 0=no	0.440		0.185		0.136		0.050	
Spouse age	years	62.3	7.51	68.6	7.27	69.8	5.81	75.4	5.16
Spouse educ	years	2.96	3.68	3.14	3.73	4.73	4.42	5.02	4.49
Spouse work	1=yes, 0=no	0.220		0.076		0.331		0.133	
Divided property	1=yes, 0=no	0.234		0.450		0.320		0.555	
Property	1=yes, 0=no	0.736		0.667		0.533		0.449	
Pension	1=yes, 0=no	0.655		0.722		0.489		0.578	
Any sons	1=yes, 0=no	0.943		0.946		0.945		0.941	
Married sons	#	1.84	1.46	2.02	1.38	2.14	1.40	2.24	1.42
Married daughters	#	2.00	1.55	2.20	1.56	2.28	1.53	2.26	1.48
Unmarried children	#	1.00	1.20	0.61	0.85	0.71	1.01	0.55	0.88

Note: means for spouse variables include married sample only.

Table 4
Age and Cohort Differences in Elderly Characteristics*

	ADL Limits	Married	Work	Property	Divided property	Children	Umarr. children	Educ	Pension
Men									
Age groups									
60-64	0.13	0.85	0.56	0.76	0.17	4.24	1.31	6.45	0.66
65-69	0.20	0.82	0.36	0.74	0.28		0.94		
70-74	0.22	0.76	0.23	0.70	0.40		0.74		
75-79	0.37	0.72	0.13	0.63	0.54		0.55		
80-84	0.62	0.66	0.04	0.62	0.69		0.51		
85-89	0.81	0.46	0.04	0.56	0.76		0.51		
Birth year cohorts									
1920-24	0.00	-0.01	0.00	-0.01	-0.07	0.47	0.09	-0.87	-0.02
1915-19	0.12	-0.04	0.03	-0.05	-0.09	0.85	-0.01	-1.64	-0.12
1910-14	0.07	-0.05	0.01	-0.01	-0.18	0.87	0.00	-2.74	-0.24
1905-09	0.06	-0.23	0.03	-0.19	-0.24	1.06	-0.09	-2.42	-0.26
<1905	0.18	-0.21	-0.01	-0.25	-0.28	0.36	-0.15	-3.88	-0.42
Women									
Age groups									
60-64	0.24	0.69	0.22	0.68	0.24	5.12	0.98	2.60	0.52
65-69	0.34	0.60	0.13	0.59	0.39		0.68		
70-74	0.45	0.51	0.08	0.55	0.50		0.51		
75-79	0.60	0.38	0.03	0.45	0.62		0.50		
80-84	0.88	0.29	0.02	0.38	0.71		0.58		
85-89	1.34	0.25	0.02	0.35	0.94		0.82		
Birth year cohorts									
1920-24	0.03	-0.07	-0.02	-0.10	-0.06	-0.02	0.05	-0.72	-0.07
1915-19	0.14	-0.09	-0.01	-0.11	-0.10	-0.08	-0.03	-1.46	-0.17
1910-14	0.15	-0.11	-0.01	-0.20	-0.10	-0.16	-0.14	-1.69	-0.24
1905-09	0.16	-0.20	0.00	-0.19	-0.29	-0.51	-0.22	-2.07	-0.35
<1905	0.10	-0.17	-0.02	-0.30	-0.37	-1.10	-0.56	-2.35	-0.33

*Numbers are coefficients of regressions on age and cohort group dummies, excluding intercepts, for characteristics that vary over time, and on a constant and cohort dummies only for time-invariatnt characteristics. The reference group is the cohort born 1925-29, aged 60-64.

Table 5
Sample Means by Living Arrangement, Pooled Data

Variable	Independent Mean	Independent Std. Dev.	With children Mean	With children Std. Dev.	With others Mean	With others Std. Dev.
Men	N=1097		N=2479		N=222	
ADL limitations	0.191	0.441	0.243	0.522	0.252	0.512
Married	0.781		0.761		0.725	
Age	70.5	5.9	70.5	6.4	70.8	7.1
Education	6.20	4.84	5.18	4.41	5.46	4.95
Work	0.329		0.297		0.396	
Spouse age	65.9	7.79	65.0	7.94	64.7	6.45
Spouse educ	3.44	3.99	2.84	3.52	3.36	4.03
Spouse work	0.180		0.169		0.149	
Divided property	0.294		0.373		0.243	
Property	0.722		0.689		0.635	
Pension	0.738		0.646		0.671	
Any sons	0.909		0.968		0.860	
Married sons	1.95	1.44	1.94	1.42	1.67	1.35
Married daughters	2.10	1.58	2.09	1.54	1.93	1.52
Unmarried children	0.55	0.95	0.93	1.09	0.57	0.84
Women	N=615		N=2564		N=230	
ADL limitations	0.365	0.488	0.516	0.673	0.621	0.752
Married	0.638		0.416		0.387	
Age	70.5	6.0	71.2	6.5	73.5	7.3
Education	2.46	3.89	1.58	2.94	1.74	3.13
Work	0.167		0.075		0.104	
Spouse age	72.1	5.54	72.3	6.13	73.2	6.15
Spouse educ	5.89	4.70	4.50	4.28	3.69	4.09
Spouse work	0.288		0.248		0.337	
Divided property	0.393		0.467		0.335	
Property	0.644		0.427		0.374	
Pension	0.615		0.489		0.470	
Any sons	0.950		0.961		0.722	
Married sons	2.36	1.41	2.19	1.38	1.57	1.47
Married daughters	2.32	1.53	2.27	1.50	2.13	1.61
Unmarried children	0.344	0.676	0.697	0.965	0.491	0.803

Note: means for spouse variables include married sample only.

Table 6
Estimation Results:
Dynamic Multinomial Probit (Full Model*)

	Men				Women			
	Independent/ With Children		With Others/ With Children		Independent/ With Children		With Others/ With Children	
Variable	Coef.	t-stat	Coef.	t-stat	Coef.	t-stat	Coef.	t-stat
constant	-0.396	-0.655	-0.666	-0.644	-2.162	-2.932	-1.436	-1.725
Independent (t-1)	1.480	9.949	0.252	1.427	1.662	11.283	0.511	2.693
With others (t-1)	0.607	2.307	1.045	3.341	0.738	2.541	1.132	4.359
ADL limits	-0.106	-1.302	0.063	0.589	-0.132	-1.496	0.093	0.944
ADL limits (t-1)	0.058	0.440	-0.062	-0.311	-0.025	-0.214	-0.129	-0.768
Married	-0.127	-0.272	0.562	0.647	0.367	0.382	0.456	0.309
Married (t-1)	-0.122	-0.903	-0.020	-0.093	-0.189	-1.264	0.139	0.821
Age	-0.310	-0.399	0.085	0.067	1.036	1.123	1.225	1.119
Education	2.689	2.665	0.210	0.147	1.263	0.933	0.695	0.382
Work	0.068	0.859	0.228	1.786	0.243	1.592	0.232	1.111
Spouse age	0.441	0.666	-0.859	-0.690	0.019	0.015	-0.784	-0.413
Spouse educ	1.461	1.135	0.816	0.442	2.741	1.878	-0.433	-0.189
Spouse work	0.198	1.787	-0.038	-0.234	0.122	0.977	0.183	0.879
Pension	0.193	2.431	0.050	0.417	0.039	0.481	0.045	0.407
Property	-0.012	-0.166	-0.238	-1.947	0.286	3.094	-0.153	-1.142
Divided property	-0.111	-1.494	-0.370	-3.025	-0.026	-0.311	-0.304	-2.414
Any sons	-0.653	-3.535	-0.722	-2.898	-0.136	-0.685	-1.087	-4.065
Married sons	0.027	0.953	-0.049	-1.012	0.040	1.291	-0.081	-1.573
Married daught.	0.008	0.375	-0.010	-0.273	0.031	1.111	-0.005	-0.141
Unmarried kids	-0.181	-4.539	-0.204	-3.007	-0.240	-3.985	-0.062	-0.820
1996	-0.114	-1.600	0.271	2.464	-0.068	-0.839	0.400	3.091
Log Likelihood		-2312				-1906		
N		2716				2530		

*Includes correlated random effects and endogenous initial conditions. Note: units for age and education variables are years/100.

Table 7
Estimated Error Correlations in Full Model

	Men		Women	
	Estimate	t-stat	Estimate	t-stat
$Var(\gamma_i^d)$	0.918	1.687	0.001	0.000
$Var(\gamma_i^o)$	1.072	2.424	0.012	0.000
$\rho(\gamma_i^d,\gamma_i^o)$	0.030	0.768	0.482	0.003
$Var(v^d_{i0})$	1.035	0.247	2.102	3.082
$Var(v^o_{i0})$	1.000	1.000	1.000	1.000
$\rho(v^d_{i0},v^o_{i0})$	0.043	0.475	0.713	0.149
$\rho(u^d_{i1},v^d_{i0})$	0.390	1.511	0.310	2.216
$\rho(u^d_{i1},v^o_{i0})$	0.000	1.492	0.158	1.659
$\rho(u^o_{i1},v^d_{i0})$	-0.069	-0.649	0.067	1.441
$\rho(u^o_{i1},v^o_{i0})$	-0.272	-1.780	-0.104	-3.111
$\rho(u^d_{i2},v^d_{i0})$	0.492	1.984	0.033	1.897
$\rho(u^d_{i2},v^o_{i0})$	0.159	1.647	0.018	1.760
$\rho(u^o_{i2},v^d_{i0})$	0.069	1.161	-0.344	-0.369
$\rho(u^o_{i2},v^o_{i0})$	-0.122	-2.771	-0.172	-2.571

Table 8
Marginal Probabilities for Full Model

	Male		Female	
Variable	Independent/ with children	With others/ with children	Independent/ with children	With others/ with children
Constant	-0.1109	0.0121	-0.3044	0.0172
Independent (t-1)	0.5059	0.0049	0.4953	0.0348
With others (t-1)	0.1385	0.1641	0.1559	0.1777
ADL limits	-0.0298	-0.0011	-0.0185	-0.0011
ADL limits (t-1)	0.0162	0.0011	-0.0035	0.0015
Married	-0.0556	0.0453	0.0783	0.0382
Married (t-1)	-0.0406	-0.0013	-0.0434	0.0093
Age	-0.0869	-0.0015	0.1458	-0.0147
Education	0.7525	-0.0038	0.1778	-0.0083
Work	0.0158	0.0227	0.0556	0.0208
Spouse age	0.1234	0.0157	0.0026	0.0094
Spouse educ	0.4088	-0.0149	0.3859	0.0052
Spouse work	0.0694	-0.0043	0.0258	0.0156
Pension	0.0613	0.0039	0.0082	0.0034
Property	0.0026	-0.0226	0.0673	-0.0098
Divided property	-0.0277	-0.0308	-0.0023	-0.0217
Any sons	-0.1827	0.0132	-0.0191	0.0130
Married sons	0.0099	-0.005	0.0092	-0.0069
Married daught.	0.0030	-0.0009	0.0065	-0.0002
Unmarried kids	-0.0547	-0.0185	-0.0524	-0.0054
1996	-0.0446	0.0263	-0.0198	0.0310

Table 9
Share of Elderly Over 65 Living Independently, 1989 and 1996

	1989		1996	
	Men	Women	Men	Women
Total	0.276	0.154	0.299	0.226
Age 65-71	0.286	0.161	0.290	0.272
Age 72+	0.260	0.147	0.311	0.181
Percent age 72+	39.8	47.1	42.9	50.9
Mean age of those 72+	76.9	76.9	77.0	77.5
Sample	1381	1186	1667	1296

*Sample aged 65-66 in 1996 from independent survey of 50-66 year olds, incorporated using appropriate sample weights.

Table 10
Simulation Results Holding Time-Varying Factors Constant at 1989 Levels

	Men			Women		
	Indep.	W/kids	W/oth.	Indep.	W/kids	W/oth.
1989 Actual Distribution	22.83	73.05	4.12	16.84	77.47	5.67
1996 Actual Distribution	28.93	63.49	7.58	21.88	69.09	9.03
1996 Baseline Prediction*	23.84	73.51	2.64	17.79	77.41	4.80
1996 Predictions Fixing 1989 Values*						
1) ADL limits	24.36	73.13	2.51	18.61	76.87	4.51
2) Marital Status	79.00	20.41	0.59	46.07	51.08	2.85
3) Age	24.26	73.15	2.59	17.05	78.41	4.54
4) Work	24.01	72.90	3.09	18.25	76.80	4.92
5) Property	23.94	73.54	2.52	18.74	76.74	4.51
6) Divided property	24.09	72.75	3.16	17.57	76.96	5.47
8) Married sons	23.60	73.66	2.73	17.63	77.57	4.80
9) Married daught.	23.79	73.54	2.66	17.74	77.47	4.80
10) Unmarried kids	22.26	75.40	2.34	16.76	78.42	4.82
11) All variables	50.81	47.48	1.71	44.99	52.11	2.90
1996 Predictions Assuming Universal Pensions	25.62	71.82	2.57	18.10	76.95	4.95

Table A1
Estimation Results:
Dynamic Multinomial Probit (IIA)

	Men				Women			
	Independent/ With Children		With Others/ With Children		Independent/ With Children		With Others/ With Children	
Variable	Coef.	t-stat	Coef.	t-stat	Coef.	t-stat	Coef.	t-stat
constant	-0.589	-0.933	-0.407	-0.389	-0.804	-1.182	-0.452	-0.560
Independent (t-1)	1.735	26.167	0.367	2.934	1.823	22.729	0.649	4.945
With others (t-1)	0.729	4.496	1.411	8.939	0.755	4.400	1.688	10.838
ADL limits	-0.113	-1.343	0.062	0.578	-0.130	-1.517	0.147	1.599
ADL limits (t-1)	0.018	0.131	-0.075	-0.388	-0.013	-0.116	-0.155	-0.936
Married	0.136	0.284	0.012	0.014	0.281	0.301	0.079	0.054
Married (t-1)	-0.135	-0.980	-0.044	-0.201	-0.266	-1.863	0.078	0.468
Age	-0.135	-0.166	-0.338	-0.263	-0.598	-0.678	-0.193	-0.182
Education	0.581	0.604	0.136	0.094	0.421	0.306	0.377	0.206
Work	0.071	0.883	0.209	1.669	0.209	1.400	0.222	1.073
Spouse age	0.127	0.187	0.000	0.000	0.246	0.200	-0.298	-0.156
Spouse educ	0.972	0.731	0.626	0.336	1.347	0.924	0.134	0.062
Spouse work	0.180	1.655	0.005	0.029	0.112	0.909	0.190	0.920
Pension	0.227	2.821	0.032	0.270	0.032	0.399	0.044	0.392
Property	0.018	0.246	-0.258	-2.173	0.266	3.103	-0.187	-1.437
Divided property	-0.101	-1.375	-0.390	-3.684	-0.041	-0.499	-0.289	-2.326
Any sons	-0.561	-3.208	-0.658	-2.838	-0.247	-1.291	-1.060	-5.270
Married sons	0.014	0.483	-0.055	-1.131	0.037	1.226	-0.075	-1.512
Married daught.	0.003	0.143	-0.016	-0.434	0.016	0.568	-0.014	-0.400
Unmarried kids	-0.178	-4.952	-0.181	-2.783	-0.230	-4.456	-0.063	-0.875
1996	-0.109	-1.607	0.265	2.533	-0.034	-0.435	0.402	3.354
Log Likelihood	-2359				-1937			
N	2716				2530			

Note: units for age and education variables are years/100.

Table A2
Estimation Results:
Multinomial Probit (No State Dependence)

	Men				Women			
	Independent/ With Children		With Others/ With Children		Independent/ With Children		With Others/ With Children	
Variable	Coef.	t-stat	Coef.	t-stat	Coef.	t-stat	Coef.	t-stat
constant	-0.668	-0.547	-0.488	-0.438	-4.163	-2.964	-2.788	-2.776
ADL limits	-0.163	-1.195	0.055	0.456	-0.312	-2.203	0.053	0.452
ADL limits (t-1)	-0.037	-0.180	-0.084	-0.372	-0.137	-0.777	-0.182	-1.018
Married	-1.830	-1.975	0.501	0.522	0.134	0.069	0.320	0.201
Married (t-1)	-0.147	-0.657	-0.133	-0.549	-0.060	-0.254	0.215	1.132
Age	0.652	0.418	0.467	0.330	2.671	1.499	3.522	2.726
Education	4.503	2.261	0.053	0.032	2.684	0.969	2.083	1.082
Work	0.258	1.745	0.306	2.173	0.752	3.176	0.467	2.031
Spouse age	2.972	2.245	-0.708	-0.514	0.687	0.270	-0.368	-0.177
Spouse educ	2.277	0.833	1.394	0.680	4.758	1.561	-2.385	-0.984
Spouse work	0.332	1.657	0.076	0.404	0.277	1.277	0.335	1.329
Pension	0.533	3.103	0.081	0.593	0.352	2.086	0.050	0.384
Property	0.028	0.213	-0.233	-1.721	0.558	3.477	-0.190	-1.298
Divided property	-0.385	-2.646	-0.482	-3.657	-0.102	-0.627	-0.484	-3.620
Any sons	-1.370	-3.889	-1.045	-3.861	-0.246	-0.667	-1.378	-5.254
Married sons	0.006	0.097	-0.078	-1.505	0.087	1.404	-0.101	-1.821
Married daught.	0.004	0.073	-0.035	-0.852	0.025	0.462	-0.012	-0.312
Unmarried kids	-0.470	-7.605	-0.309	-4.177	-0.595	-5.644	-0.192	-2.160
1996	0.013	0.136	0.342	2.893	0.177	1.566	0.442	3.293
		Coef.	t-stat			Coef.	t-stat	
Var(γ_i^d)		4.076	112.690			3.863	86.066	
Var(γ_i^o)		0.616	1.152			0.633	1.024	
Cov(γ_i^d,γ_i^o)		0.541	-0.821			0.510	-0.937	
Log Likelihood		-1692				-1373		
N		2716				2530		

*Includes correlated random effects. Note: units for age and education variables are years/100.