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# Population aging and the extended family in Taiwan: a new model for analyzing and projecting living arrangements

**Andrew Mason**

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# Population aging and the extended family in Taiwan: a new model for analyzing and projecting living arrangements

**Andrew Mason**

**Sang-Hyop Lee**[1]

## Abstract

Population aging produces changes in the availability of kin with uncertain implications for extended living arrangements. We propose a highly stylized model that can be used to analyze and project age-specific proportions of adults living in extended and nuclear households. The model is applied to Taiwan using annual data from 1978-1998. We estimate cohort and age effects showing that more recently born cohorts of seniors are less likely to live in extended households, but that as seniors age the proportion living in extended households increases. The effect of individual aging has diminished over time, however. The proportion of non-senior adults living in extended households has increased steadily because changes in the age structure have increased the availability of older kin. The model is used to project living arrangements and we conclude that the proportion living in extended households will begin to decline gradually for both seniors and non-seniors. The extended family is becoming less important in Taiwan, but it is not on the way out.

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## 1. Introduction

Living arrangements in Asia are quite distinctive as compared with those found in Western societies. The great majority of Asia's elderly live with their adult children. Living arrangements are beginning to change, however, and in some countries quite rapidly. In some East Asian countries the proportion of seniors living with their children has dropped substantially in recent years. The situation in Taiwan is somewhat distinctive. The proportion of seniors living with their children has been relatively stable in recent years and the proportion of non-senior adults living in extended households has increased. As the analysis presented below will show, however, recent aggregate trends masks important underlying changes that will lead to a steady decline in extended living arrangements. Further decline is on the horizon, but the evidence does not point to a precipitous decline. At least in Taiwan, extended living arrangements do not appear to be on their way out.

To many, the decline of the extended family is a matter of concern particularly when many Asian countries face rapid population aging. The extended family has been the primary system of support for the elderly. Their children have provided both financial support and personal care to their aging parents. Some Asian governments have gone so far as to mandate that children support their elderly parents.

Others have noted, however, that the traditional family support system may be ill-suited to a society with large numbers of dependent elderly. Middle-aged women, in particular, may face a heavy burden as the sole care-giver for their parents, the parents of their husband, and in some cases grandparents. Alternative support systems can spread the costs of caring for the elderly. The family support system is also an unattractive way of financing retirement in an aging society for exactly the same reason as PAYGO pension programs. In slowly growing or declining populations, the rate of return that can be sustained from intergenerational transfers is much lower than the rate of return available from financial markets or real investment (Lee, Mason et al. 2000).

The objective of this paper, however, is not to address whether the decline in the extended family is a welcome development or not. The purpose is to model a specific feature of living arrangements – the choice by families to form multi-generation extended households – and to make some assessment about the future course of extended living arrangements. The motivation is an interest in how population aging is influencing familial support systems for seniors. The model that we propose abstracts from other social changes that influence living arrangements. For example, our model does not distinguish nuclear households from joint households consisting of married siblings, but adults who are only members of one generation. Nor do we consider decisions by the young regarding the transition to adulthood even though this will indirectly influence living arrangements. For example, if children delay the age at which they are leaving school, marrying, and establishing an independent residence, this will influence living arrangements. But these decisions may have more to do with what it means to be an adult and to achieve economic independence from one's parents and less to do with how familial support systems respond to the needs of seniors. The advantage of our approach is that it focuses attention on a key issue – how population aging is influencing multi-generation extended living arrangements. The drawback is that it neglects important aspects of living arrangements that are treated in more comprehensive models (Zeng et al., 1997, 1998).

The approach employed here is macro-analytic and has more in common with other macro approaches than with the microsimulation methods pioneered by Gene Hammel, Ken Wachter and their collaborators (Hammel et al., 1981, 1991). Some of the important underlying issues – the availability of kin, for example – are much more readily addressed through microsimulation approaches, although the implications for macrosimulation are obvious and important. Lin's (1994) application of SOCSIM is of interest, for example, because he provides simulated values for the proportion of Chinese elderly with surviving children.

Our approach is similar in some respects to household projection research (Kono, 1987; Pitkin and Masnick, 1987; Holmberg, 1987; Mason, Racelis et al., 1996; Zeng et al., 1997, 1998) that attempts to devise a model that can be used to project or forecast a key aggregate characteristic of households, e.g., a headship rate. The extent to which the underlying processes are analyzed varies among these models, but they do not rival the detailed treatment found in SOCSIM and other microsimulation methods. The model analyzed in this paper abstracts from many of the details of important demographic processes, but does not neglect them altogether.

We are aware of a few efforts to project the proportion of individuals living in extended households. Freedman et al. (1991) projects the percentage of US women 85 and older living with children and confronts similar issues to the ones addressed in this study. Zeng et al. (1997, 1998) have developed a very detailed household projection model with compositional detail that is sufficient to allow the calculation of the age-specific proportions of children and non-senior adults living in three-generation households and the proportion of elderly (total, but not age-specific) living in three-generation households. Other household projection models often distinguish different types of households or the age and sex of household members, but it is difficult to infer characteristics of individuals from the set of characteristics of households that are typically available from household projections. For many purposes, such as a description of the extent of family support available to seniors, the individual (the senior) is the analytic unit of interest, not the household.

In this paper we analyze changes in living arrangements by seniors and by non-senior adults. The two are connected, of course, because extended households as we have defined them here are created when seniors and their adult children choose to live together. Thus, an increase in the proportion of seniors leads to an increase in the proportion of working-age adults (non-seniors) living in extended households, *ceteris paribus*. The relationship is not an entirely simple one, however. The proportion of working-age adults living in extended households will depend on the extent to which the "dependency burden" that arises from co-residence is equally shared across the population. To illustrate the point one could imagine two highly unrealistic extremes. All seniors in a population could live in a single household with only one non-senior adult present. In this case, the proportion of seniors in extended households would be one and the proportion of non-seniors living in extended households would be essentially zero because only one non-senior adult would live in an extended household. The old-age dependency ratio within extended households would be extremely high. At the other extreme, an entirely egalitarian world, seniors would spread themselves evenly among all non-senior members of the population. In this case, all seniors and all non-seniors would live in extended households and the old-age dependency ratio in extended households would be identical to the old-age dependency ratio for the population at large. The key point here is that the connection between the proportion of seniors and non-seniors living in extended households is not a mechanistic one. Rather it is a flexible one that will differ among countries and over time depending on how the dependency burden of senior co-residents is distributed among the non-senior members of the population.

There are important demographic/social constraints that preclude the unrealistic living arrangements described above. An important one is that extended households typically consist of kin and, hence, that availability of kin influences whether or not the burden of supporting seniors is widely shared or not. Many non-senior adults may not live with elderly parents because their parents are deceased or because one of their siblings is living with their parents. Thus, given the proportion of seniors living in extended households, changes in fertility and mortality may have a large effect on the proportion of non-seniors living in extended households. Freedman et al. (1991) provide a formal treatment of the relationship between fertility, kin availability, and living arrangements.

This issue is addressed in a formal, but tractable, way below. In the empirical analysis, we find that the distribution of the dependency burden has changed over time so that the proportion of non-seniors living in extended households has increased relative to the proportion of seniors living in extended households. There are some indications, however, that this phenomenon will not continue in the future and that the proportion of seniors and non-seniors living in extended households will both decline.

## 2. Extended living arrangements: what drives the decision?

The family offers an efficient institution for producing, consuming, and redistributing resources among family members and across generations. In all societies, the family is the primary institution through which resources are transferred to children from productive members of the population (parents). As children age and become productive, resources may begin to flow from children to their parents (Lee 2000). The extent and direction of resource flows between prime-age adults and their elderly parents is perhaps more complex. Decisions by the family may be governed by altruism (Becker 1981). The extended family may offer an efficient institution for exchange. Elderly parents may care for their grandchildren in exchange for financial support. Or, adult children may care for the elderly parents in return for a bequest. Or, the family may provide a system by which family members pool risks that they face in various aspects of their daily lives (Kotlikoff and Spivak 1981). The extended family may offer scale economies in consumption that allows members to achieve a higher standard of living (McElroy and Horney 1981). Of course, the roles are not mutually exclusive nor is the list provided here exhaustive.

Fulfilling many of these functions does not require co-residence. Family members can live in independent households and exchange goods and services (money or time), but co-residence provides an efficient means for carrying out these transactions. If the transactions are large and frequent and if they involve time, family members may choose to live together throughout their lives. If the transfers tend to be episodic or confined to a particular period during the lifecycle, family members may choose to vary their living arrangements depending on the current circumstances.

One of the empirical features of living arrangements in Taiwan is that for many seniors co-residence is not a permanent feature. The proportion of seniors living with their adult children increases with the age of those seniors. In the absence of panel data, however, we cannot tell whether the decision by seniors to live with their children is a permanent one or whether it is common for seniors to change living arrangements frequently. Some seniors may rotate their residence from one child to another so that they are permanently in an extended household, but their children are not. In any event, the rise among seniors in extended living arrangements as they age is consistent with the view that the formation of extended families is a response to changing circumstances associated with aging, i.e., the risks associated with growing old.

The risks faced by the elderly take a number of forms. Older adults face substantial financial risks. Three that directly affect financial wealth seem particularly important: premature forced retirement, investment risk due to fluctuations in financial markets, and longevity risk (the risk of living longer than expected and, hence, having insufficient resources for the later years of life). Older adults also face risks on the consumption side of which unanticipated health care expenditure looms large.

As older adults age, they experience a succession of shocks that reduce their financial resources, and for some, to unexpectedly low levels. The financial hardship they experience may induce transfers from their children and ultimately lead to co-residence (Note 1). In the absence of uncertainty, there is no obvious reason why the economic situation of the elderly would decline as they age, so long as the elderly were effective lifecycle planners.

Financial risk is not the only uncertainty faced by the elderly. Indeed, it may not be the most important factor that leads to increased co-residence. The elderly face risks that may also greatly affect the value they attach to personal attention or time inputs from their children. Two aspects seem particularly important – the first is health risks. The elderly face health crises that influence the need for personal care. Because personal care from strangers is a poor substitute for personal care from children and because personal care requires direct contact, the efficiency gains from co-residence may be especially large. The second event faced by the elderly that influences their demand for attention from their children is the death of their spouse. A spouse may be a source of companionship and a source of personal care. The loss then will lead to increased demand for companionship and personal care from children. In addition, the loss of a spouse may also influence other calculations of the cost and benefits of co-residence. The economies to be gained may be substantially larger when a single-person household is absorbed rather than a two-person household.

The family is not the only means of insuring against these risks. Both commercial and social insurance can play an important role. Wealth can be annuitized by participating in employer-based defined benefit retirement programs, by purchasing commercial annuities, or by participating in public pension programs. These programs protect the elderly from both investment risk and longevity risk. Health insurance, either private insurance or the public provision of health care, reduces the financial risks associated with illness. Moreover, individuals can self-insure by accumulating more wealth during their working years.

The availability of risk-sharing alternatives to the family increases as societies develop and as growth in the number of elderly increases the demand for commercial and social insurance. Pension programs develop and are extended to growing numbers of workers employed in the public sector and by larger private firms. Comprehensive health insurance becomes increasingly available and may be extended to retirees. Publicly funded health programs meet the health care needs of the elderly. But even in the most economically advanced economies, the elderly face risks that are difficult or impossible to insure. The market for annuities is thin and the price is so high that few elderly can protect themselves against investment and longevity risk in the absence of comprehensive social insurance (Note 2). Many of the other risks described are essentially uninsurable. Thus, there is little reason to think that the role of the family, and intergenerational co-residence, will shrink to nothing as Asian societies develop.

The theory of co-residence implicates a number of factors that may account for changes in the extent to which adult children and their parents co-reside. The effects of kin availability, per capita income, education, and disability have all been explored in empirical studies (Bachrach, 1980; Chevan and Korson, 1975; Kobrin 1976; Macunovich et al., 1995; Michael, Fuchs, and Scott, 1980; Soldo, 1981; Wister and Burch, 1983; and Wolf, 1995). The development of social and commercial insurance may lead to a decline in the importance of the family support system and to a decline in co-residence. But as will be seen below, the analysis presented here does not attempt to identify which specific aspects of social and economic development influence changes in living arrangements.

## 3. Modeling extended living arrangements

We begin with a highly stylized household model that abstracts from many of the details of living arrangements and family connections. This approach yields a tractable macro-level model, with an inevitable sacrifice of some reality.

Assume that adults in a one-sex population belong to one of two generations: persons who are at least $g$ years old but less than $2g$ years of age are called non-seniors; persons who are $2g$ years of age or older are called seniors, where $g$ is the length of a generation and no person survives beyond age $3g$ (Note 3). Children, those under age $g$, do not play a role in the model except that they provide the next generation of non-seniors.

All births are to individuals of age $g$. Given this important assumption the number of family members within a population is readily identified. Non-seniors aged $a$ have parents aged $a+g$; seniors aged $a+g$ have children who are aged $a$. $N(a)$ is the population aged $a$, $N(a+g)$ is both the population aged $a+g$ and the parents of persons aged $a$. $N(a+g)/N(a)$ is the number of surviving parents per non-senior aged $a$ (Note 4).

All persons live in one of two types of households: extended households, consisting of both non-senior adults and seniors; and, nuclear households, consisting of non-senior adult or seniors, including persons living alone, but not both. All extended households are formed from family members. Hence, extended households with a non-senior aged $a$ have only senior members aged $a+g$. The non-senior population aged $a$ living in extended households is $N^x(a),\quad g \le a < 2g$, and the senior population living in those same extended households is $N^x(a+g)$.

Let $n(a,t)$ for be the proportion of persons aged $a$ living in nuclear households in year $t$ and $x(a,t)$ be the proportion living in extended households. The proportion of seniors $(a \geq 2g)$ living in a nuclear household depends on a cohort effect, $k(b)$, where $b = t - a$ is the year of birth, and an age effect, $h(a)$. The cohort effect captures persistent, lifetime characteristics of members of any birth cohort and the age effect captures the influences of changes that each birth cohort experiences as it ages: (Note 5)

$$n(a,t) = 1 - x(a,t) = k(b)h(a) \quad \text{where } a \geq 2g \text{ and } b = t - a. \tag{1}$$

The proportion of non-seniors ( $g \leq a < 2g$ ) living in extended households depends on two demographic factors. First, the non-senior population provides the members that live with seniors in extended households. Thus, the proportion of non-seniors living in extended households varies directly with the proportion of seniors living in extended households.

Second, the interdependence between the proportion of seniors and non-seniors living in extended households depends on the age composition of extended households. If the age structure of extended households were identical to the age structure of the general population, then the proportions of seniors and non-seniors living in extended households would be identical. If the number of seniors per non-senior is greater in extended households than in the general population, then the proportion of non-seniors living in extended households will be less than the proportion of seniors living in extended households.

The interdependence between the proportion of seniors and non-seniors living in extended households is incorporated into the analysis using age-specific dependency ratios that approximate the age structure of families $(d(a,t))$ and extended households $(d^x(a,t))$. It is straightforward to show that the following identify holds:

$$x(a,t) \equiv ddx(a,t)x(a+g,t) \quad \text{for } g \leq a < 2g, \tag{2}$$

where:

$$\begin{aligned} &ddx(a,t) = d(a,t) / d^x(a,t), \\ &d(a,t) = N(a+g,t) / N(a,t), \text{ and} \\ &d^x(a,t) = N^x(a+g,t) / N^x(a,t) \quad \text{for } g \leq a < 2g. \end{aligned} \tag{3}$$

The age in this formulation is the age of the non-senior members of the population or the extended household. In the application below the values are constructed for five-year age groups.

In general, the relative dependency structure in the population (*ddx*) will depend on fertility, mortality, migration, living arrangements and their inter-relationships. For the highly stylized model employed here, however, we can formalize the relationship. Let *f(a+g)* represent the net reproduction rate and the *s(a+g)* the proportion surviving from age *a* to age *a+g* for the population currently aged *a+g*. The dependency ratio for the general population aged *a* is *d(a,t)=s(a+g,t)/f(a+g,t).* The dependency ratio for extended households will vary depending in part on the form of extended households. At one extreme seniors may live with all of their surviving adult descendants. In this case, $d^x(a,t) = 1/f(a+g,t)$. At the other extreme seniors may live with only one of their surviving adult descendents (Note 6). In this case, $d^x(a,t) = 1$. Note that in the special case of replacement fertility, the dependency ratio in extended households would be the same given either form of living arrangements. For the stylized model, we have:

$$
\begin{aligned}
ddx(a,t) &= s(a+g,t)/f(a+g,t) && \text{for stem extended households, and} \\
&= s(a+g,t) && \text{for laterally extended households.}
\end{aligned}
\qquad (4)
$$

If equation (4) is an adequate characterization of the relative dependency structure, we can expect the value of *ddx* in a society characterized by stem households to be close to 1 for *a*=30, to decline with *a* reflecting the underlying survival schedule, and to increase with *t* due to improving survival chances. If laterally extended households are common then values of *ddx* will be lower so long as fertility is above replacement level, but *ddx* will rise more rapidly over time if the net reproduction rate is declining.

There are additional factors not captured by our highly stylized model that should influence *ddx*. First, equation (4) is based on the assumption that survival chances are independent of living arrangements. However, living arrangements may influence survival rates either because one household form provides a healthier environment or because health status influences choice of living arrangements. Mortality may also influence the relative dependency structure because of its effect on living arrangements. The first death of a parent may increase the chances that the surviving parent will move into an extended household. Such an event will generally reduce the dependency ratio within extended households by establishing a new household with only one senior member. The dependency ratio of the general population will also be reduced, but it is unclear whether the decline is necessarily smaller or greater than the decline in the dependency ratio in the extended household population. The second death of a parent has a very clear effect. The dependency ratio in the general population will decline. If the parent lived in an extended household, the household will be transformed from extended to nuclear. This will lead to a rise in the average dependency ratio of the remaining extended households because extended households that lose their last senior member have low dependency ratios relative to the mean.

Second, the net reproduction rate of seniors (*f*) living in extended households may differ from the net reproduction rate of seniors living in nuclear households. Of course, to live in an extended household requires at least one surviving adult descendant (Note 7). Those with more surviving children are more likely to have a child that is a suitable candidate for co-residence (rich, generous, and in Taiwan male). Moreover, childbearing may be influenced by the taste for or expectations about living arrangements. If so, those who expect to live in extended households may choose to bear more children. On the other hand, seniors with lower fertility rates may have invested more per child influencing survival chances for their children or other factors that affect the likelihood of co-residence.

Third, as presented above equation (4) treats a population as following one of two forms of extended living arrangements where in reality most societies will be a mixture of these two extremes. Moreover, the form of living arrangements may be changing over time or may vary by age. In East Asia, the stem family has become the norm although some seniors live with more than one of their surviving offspring either 'permanently' or during a period of a year or two after their children first marry (Note 8). Thus, *ddx* may be lower than the survival rate for low values of *a* even in societies where the stem family system is dominant.

In the empirical analysis presented below we explore how the relative dependency structure varies with age and across cohorts. Combined with analysis of the proportion of seniors living in extended households, we estimate a new model for projecting the proportion of seniors and non-seniors living in extended households based on equations (1) and (2).

## 4. Taiwan – is the extended family on the way out?

In some respects Taiwan is an ideal subject for this research. Demographic and economic developments have been so rapid that the twenty-one years of data at our disposal covers a period of enormous change. In the early 1950s, Taiwan had barely begun its demographic transition and its people were quite poor. By 1999, per capita GNP had reached $13,250, life expectancy at birth was 78 for females and 72 for males and the total fertility rate was 1.6 births per woman. The population is beginning to age. In 2000 an estimated 8.6 percent of the population was 65 and older. Income is very equally distributed – the Gini coefficient is 0.33. Its levels of educational attainment are very high – gross enrollment ratios for secondary school are 101 for females and 98 for males. About 60 percent of the population lives in urban areas (ADB 2001). Thus, in five decades Taiwan has transformed itself into one of Asia’s most advanced economies. Only Japan, Hong Kong, and Singapore can boast of a higher standard of living.

One idiosyncratic feature of Taiwan’s demography is relevant to the analysis. Around 1950, a large group of mainland Chinese migrated to Taiwan. The group was about 15 percent of Taiwan’s population at the time. About two-thirds were men and they were heavily concentrated in the young adult ages. The great majority were not accompanied by their parents and, hence, could not establish extended households with their parents. Moreover, many never married nor bore any children. As a consequence they cannot form extended households by living with children.

How rapidly living arrangements are changing in Taiwan is an open question. Some data suggest that extended living arrangements will persist. More than half of those 60 and older live in a multi-generation extended family. Of those 85 and older in 1998, three-quarters lived in a multi-generation extended family. The proportion 60 and older living in an extended household changed very little between 1978 and 1998 – varying between 57 percent and 52 percent with no clear trend. During the same period, the proportion of those aged 30-59 living in extended households increased from under 20 percent to nearly 30 percent.

Other data point to a decline in extended living arrangements. Weinstein et al. (1994) report that the percentage of married couples aged 20-39 who reported living with their parents at least one month after marriage declined from 90 percent in 1967 to 70 percent in 1986. They also report that the percentage of parents with a married son who lived with a married son declined from 82 percent in 1973 to 70 percent in 1986.

As we shall see the apparent stability in extended living arrangements masks important changes that will eventually lead to a decline in extended living arrangements. The extended family is not on the way out in Taiwan, but it will become less important over time.

## 5. Data

We use the Survey of Family Income and Expenditure in Taiwan (FIES), also known as the Survey of Personal Income Distribution in Taiwan until 1993. The FIES was first conducted in 1964 and, then, every other year until 1970. Since 1970 the survey has been conducted annually and data are available for the 1976 and subsequent surveys.

For technical reasons, we have confined our analysis to surveys conducted in 1978 until 1998. The number of households surveyed has varied over time, but the sample size is more than sufficient for our purposes. In 1998, about 0.4 percent of all households (14,031 households and 52,610 individuals) were covered. These are not panel data, but repeated cross-sections.

The living arrangements model is estimated using mean values of the proportions living in nuclear households, extended households, the dependency ratio, the dependency ratio in extended households, and other variables by current age (single years of age; 85 and older is the upper bracket) and birth cohort. The senior population consists of all those who are 60 or older (Note 9). The data set yields 546 observations, 21 years times 26 age groups. The oldest birth cohort was born in 1893 and the youngest birth cohort was born in 1938. For these two birth cohorts we have values for only one age – 85+ for the oldest cohort and 60 for the youngest cohort. For several birth cohorts we have values for 21 age groups, but no cohort can be followed over the entire aging period 60-85+. The non-senior population consists of all those who are 30-59. The data set yields 630 observations on non-seniors, 21 years times 30 age groups. The oldest cohort was born in 1923 and the youngest in 1968. Again we have only a single value for the youngest and oldest cohorts and values at 21 ages for those born around 1955.

Households were classified as nuclear or extended in the following manner. All related household members are classified into generations using relationship to head information. For example, the head's generation consists of all those reporting head, spouse of head, or sibling of head. The parent's generation consists of all those reporting parent, aunt, or uncle of the head. Any household that includes adult members belonging to two different generations is classified as an extended household. Note that marital status is not a factor in the classification scheme.

The cut-off for being an adult is 30 years of age. A relatively late age is used because our focus is on the support system for seniors, not the support system for young adults. To the extent possible, we do not want the trends in extended living arrangements as we measure them to be the result of a longer period of child dependency.

Selected data are plotted in Figure 1. Panel A shows the mean proportions of seniors living in nuclear households by age for selected birth cohorts. Panel B shows the mean values of *ddx* by age for selected non-senior birth cohorts (Note 10).

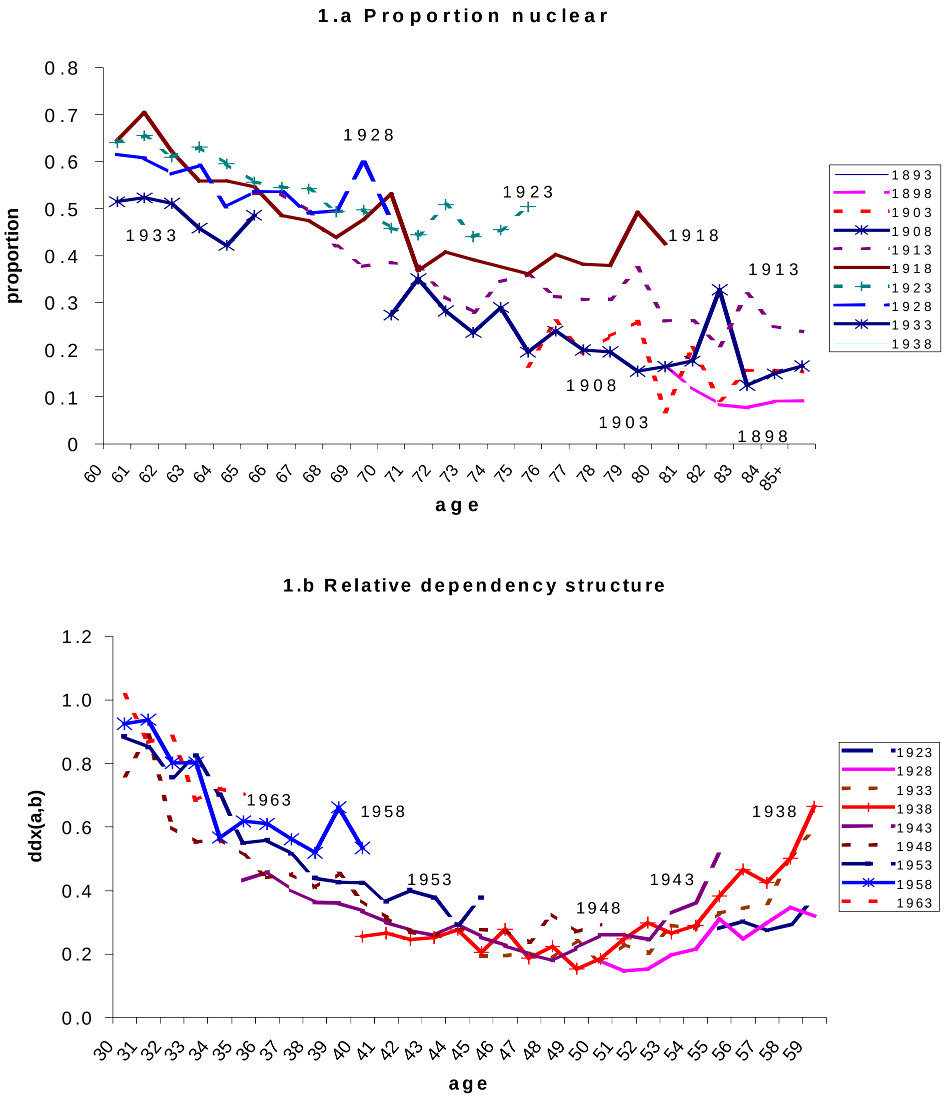


***Figure 1:*** *Proportion nuclear and the relative dependency structure by age and birth cohort*

## 6. Specification 1: Fixed age and cohort effects

In the remaining sections of this paper we explore alternative approaches to specifying the model of extended living arrangements. The first specification assumes that the proportion of seniors living in extended households and *ddx* are the product of fixed age and cohort effects. For this specification a critical assumption is that age effects have not varied across cohorts. This represents a null hypothesis in the sense that the alternative specification considered below assess whether age effects interact with cohort effects and whether alternative representations of the age effects can reduce or eliminate the interaction that is found to exist. In the absence of interactions, however, the cohort and age effects for the proportions of seniors living in nuclear households are estimated by:

$$\ln n(a,b) = \sum_{a=60}^{85+} \alpha_k A_a + \sum_{b=1893}^{1938} \beta_b B_b \tag{5}$$

where $A_a$ is a dummy variable that takes on the value of one for cohorts of age *a* and $B_b$ is a dummy variable that takes on the value of one for cohorts born in year *b*. The age effect, i.e., the coefficient $\alpha_k$, is an estimate of the natural log of the proportion of persons aged *a* living in nuclear households relative to the proportion of persons aged 60 living in nuclear households controlling for birth cohort (Note 11). The cohort effect, i.e., the coefficient $\beta_j$, is an estimate of the natural log of the proportion of 60-year-olds by year of birth *j* living in extended households.

Similarly, the relative dependency structure can be estimated by:

$$\ln ddx(a,b) = \sum_{a=30}^{55} \gamma_k A_a + \sum_{b=1919}^{1968} \eta_b B_b. \tag{6}$$

The age and cohort coefficients have similar interpretations as in the specification of the proportion living in nuclear households. The cohort effect is an estimate of the natural log of the relative dependency ratio at age 30 by birth cohort. The age effect is an estimate of the natural log of *ddx* at age *a* relative to the value at age 30 controlling for year of birth.

The estimates are presented in Figure 2. Panels A and B show the cohort and age effects, respectively, for the proportion of seniors living in nuclear households. The cohort effect rises substantially among early birth cohorts. The estimated proportion of the 1922 birth cohort living in nuclear households at age 60 exceeds 0.7 as compared to less than 0.3 for the 1900 birth cohort. The estimated proportion of more recent birth cohorts living in nuclear households is much less than 0.7. For the most recent birth cohort, those born in 1938, the estimated value is less than 0.5.

The reversal in the cohort trend towards nuclear living arrangements is a surprising phenomenon, but the high proportions nuclear may reflect the imbalance between men and women due to the large-scale immigration *circa* 1950 discussed above. The sex ratio at age 60 is a convenient measure of the impact of immigration on the marriage market. The correlation with the proportion nuclear is very evident, as shown in Figure 2, suggesting that the proportion living in nuclear households was elevated to unusually high levels by the enormous surplus of males found in the 1912-1930 birth cohorts.

The estimated age effects, the proportions nuclear at age *a* relative to the proportion nuclear at age 60 for any birth cohort, are substantial. An increase in age from 60 to 70 reduces the proportion nuclear to less than 0.7 of the age 60 level. By age 80 the proportion nuclear is roughly 0.5 times the age 60 level and by age 85 and older, 30 percent of those living in nuclear households at age 60 are still doing so.

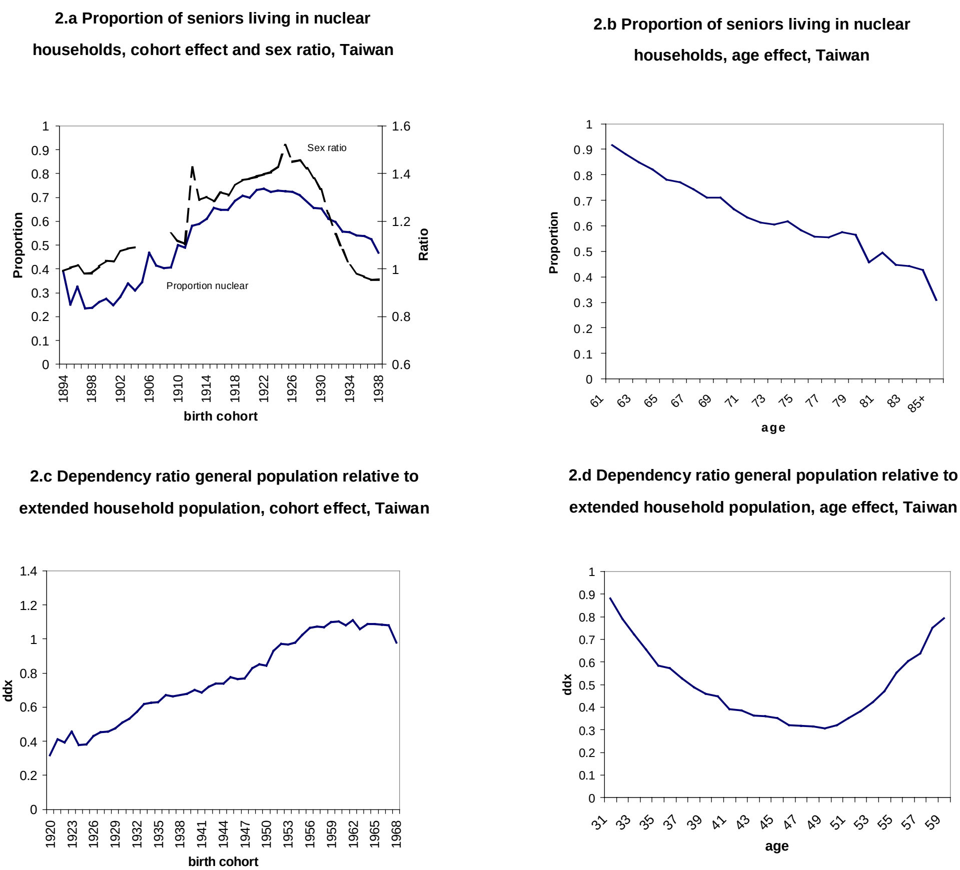


***Figure 2:*** *Estimated age and cohort effects, fixed age effects, proportion nuclear and relative dependency structures*

The relative dependency structure in Taiwan is also characterized by substantial cohort and age effects shown in panels C and D, respectively. Among early birth cohorts the *ddx(30)* estimate is only 0.4, but the value increases steadily so that by the time the 1955 birth cohort turned 30 *ddx(30)* had reach 1. Among later birth cohorts, the 1968 cohort excepted, *ddx(30)* has ranged from 1.05 to 1.1. As will be shown below, however, the cohort trend is apparently an artifact of the assumption that the age effects do not interact with the cohort effects.

The estimated age effects are large. By age 35, the value of *ddx* has declined to only 0.6 of its age 30 level. By age 50, *ddx* has declined to about one-third of its age 30 level. After age 50, the dependency ratio of the general population rises relative to the dependency ratio within extended households. The downward portion of the age effect is broadly consistent with the stylized model presented above that *ddx* declines with the survival rate. However, the rate of decline in the estimated *ddx*, especially at younger ages, is too rapid to be explained solely by declining survival.

## 7. Specification 2: Variable age effects

The assumption that the age effects don't interact with cohort effects is convenient, but seems quite implausible. There have been substantial improvements in mortality and, possibly, health (Note 12) and the development of alternative support systems (social systems provided by the government; market based systems that supply insurance and facilitate the accumulation of personal wealth) that would surely influence the effects of aging on co-residence. Likewise, the *ddx* age effects have surely changed over time in response to changes in age-specific survival rates, fertility rates, and living arrangements as discussed above.

We explore whether or not age effects have shifted, whether the changes are large and whether they are relatively systematic by specifying a functional form for the age effects and allowing the age effects to vary across birth cohorts. For the proportion nuclear we assume that $\ln h(a)$ in equation (1) is linear in age, in which case we have:

$$\ln n(a,b) = \ln n(60,b) + \beta_b (a-60). \tag{7}$$

The birth cohort effect is estimated using dummy variables for single year of birth as in specification 1. The effects of age on the proportion nuclear, $\beta_b$, are estimated for four broad birth cohorts by interacting *(a-60)* with dummy variables.

Given the non-linear, non-monotonic relationship between the dependency structure of the population and age exhibited in Figure 1, we estimate the age effect in the *ddx* equation using a quadratic:

$$\ln ddx(a,b) = \beta_{0b} + \beta_{1b}(a-30) + \beta_{2b}(a-30)^2. \tag{8}$$

The parameters have year-of-birth subscripts because we use dummy variables to estimate separate coefficients for four broad birth cohorts. The intercept term is an estimate of $\ln ddx(30,b)$ because the other two right-hand-side terms are zero for age equal to 30. Equations (7) and (8) are estimated using ordinary least squares. All estimates were obtained using the consistent variance-covariance matrix estimator of White (1980). The standard errors are thus robust to heteroscedasticity. The model is estimated using the full sample of mean values for single years of age for the available birth cohorts yielding 546 observations as for specification 1. The age variables are interacted with dummy variables that correspond to the four broad birth cohorts that are distinguished. For the proportion of seniors living in nuclear households the broad birth cohorts are: 1893-1908, 1909-1915, 1916-1922, and 1923-1938. For estimating the dependency structure the available birth cohorts (of non-seniors) are 1919-1935, 1936-1943, 1944-1951, and 1952-1968. The birth cohorts were selected so that each group contains a roughly equal number of observations.

One of the difficulties encountered in estimating age effects in the *ddx* model is that for some cohorts there are no observations over a substantial portion of the age range of interest. This would present no apparent problem were the age effects linear. A problem does arise, however, when using a flexible functional form such as a quadratic. In the empirical work the problem arises for the 1923-1938 birth cohort for which we have no observations for ages 30-39. When we fit the quadratic to this birth cohort the age effect rises very steeply at young ages in a manner quite inconsistent with the results obtained from younger birth cohorts. To deal with this problem, we estimate a second model that restricts the age effects for the 1923-38 birth cohort to equal the age effects for the 1939-45 birth cohort for the 30-39 ages using a quadratic spline. The functional form employed is:

$$\begin{aligned}\ln ddx(a,b) = {} & \beta_{0b} + B2345(\beta_1(a-40) + \beta_2(a-40)^2)\mathrm{A39+} \\ & + B3945(\beta_3(a-30) + \beta_4(a-30)^2).\end{aligned} \tag{9}$$

where *B2345* and *B3945* are dummy variables that take on the value of one for the 1923-45 and the 1939-45 birth cohorts and *A39+* is a dummy variable that takes the value of one if age is greater than 39. Using this specification, the age effect for the 1923-38 birth cohort deviates from the age effect for the 1939-45 birth cohort at age 40, but the function is continuous at 40 as is its first derivate.

The estimated cohort effects for the proportion of seniors living in nuclear households are charted in Figure 3. Comparing Figure 3 with Figure 2A shows that the estimated cohort effects are not greatly influenced by allowing for interaction between cohort and age effects. For 60-year-olds we find a rapid transition from extended to nuclear households followed by a substantial reversal. The peak is only slightly lower in Figure 2 when we allow for variable age effects.

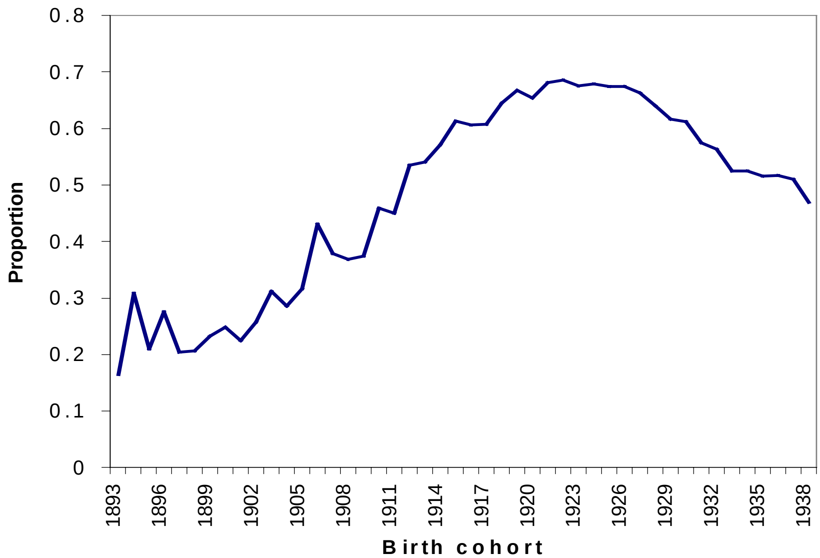


***Figure 3:*** *Estimated effect of year of birth on proportion nuclear*

Note: see Table 1 for age effect

The estimated age effects for the proportion of seniors living in nuclear households are provided in Table 1. The age effects change from early born to later born cohorts. For the 1893-1908 birth cohort the natural log of the proportion nuclear declines by –0.069 for each additional year of age. The age effects for later born cohorts are substantially smaller. Seniors belonging to more recent birth cohorts are moving in with their children at a later age than in the past.

***Table 1:*** *Estimates of age effects for proportion nuclear (60 and older)*

| **Variable & birth cohort** | **Proportion nuclear** | |
|---|---|---|
| | **Coefficient** | **Standard error** |
| Age - 60 | | |
| 1893-1908 | -0.069 | 0.012 |
| 1909-1915 | -0.032 | 0.003 |
| 1916-1922 | -0.027 | 0.002 |
| 1923-1938 | -0.024 | 0.001 |
| $R^2$ | 0.98 | |
| N | 546 | |

The cohort and age effects for *ddx* are presented in Table 2. Allowing for changing age effects yields a very different picture than the estimates based on a fixed age effect presented above. The cohort effects are much weaker than in Figure 2 and the strong upward trend prior to the mid-1950s has disappeared altogether. For three of the broad birth cohorts, the estimated cohort effect is not significantly different from zero. For the two earliest birth cohort the corresponding values of *ddx(30),* the exponential of the estimated coefficients, is about 1.2. For the youngest cohort the estimated value of *ddx(30)* is essentially 1. The only significant difference among cohorts is for the third birth cohort (1944-51) for which the estimated value of *ddx(30)* is 0.88, exp(-0.131). The strong and steady trend that is apparent in Figure 2 is apparently an artifact of the assumption of an unchanging age effect.

***Table 2**:* *Estimates of cohort and age effects for the relative dependency structure*

| Variable & birth cohort | Dependency structure | |
|---|---|---|
| | Coefficient | Standard error |
| Cohort effect | | |
| 1919-1943 | 0.2006 | 0.1331 |
| 1936-1943 (1) | 0.1793 | 0.0932 |
| 1944-1951 | -0.1297 | 0.0276 |
| 1952-1968 | -0.0012 | 0.0126 |
| Age - 30 | | |
| 1919-1943 | -0.2028 | 0.0113 |
| 1936-1943 (1) | -0.0343 | 0.0212 |
| 1944-1951 | -0.1343 | 0.0060 |
| 1952-1968 | -0.0976 | 0.0059 |
| $(\text{Age-30})^2$ | | |
| 1919-1943 | 0.0061 | 0.0003 |
| 1936-1943 (1) | 0.0007 | 0.0009 |
| 1944-1951 | 0.0038 | 0.0003 |
| 1952-1968 | 0.0018 | 0.0004 |
| $R^2$ | 0.98 | |
| N | 630 | |

Note:
1 Coefficients are the difference between 1919-35 cohort and the 1936-43 cohort for age>39.

The changes in the age effect parameters are small but systematic. The coefficient of the linear term increases from -0.203 for the 1919-35 cohort to -0.098 for the 1952-68 cohort (Note 13). The coefficient of the squared term declines from 0.0061 for earliest birth cohort to 0.0018 for the most recent birth cohort. Both the cohort and age effects are incorporated into Figure 4, which shows the predicted value of *ddx* for the four birth cohorts. Allowing for a changing age effect also leads to a very different picture than the constant age effect result. The rise in *ddx* at older ages is apparently much smaller than suggested by the fixed age effect model. For the most recent birth cohort there is no increase at all in the estimated age effects, however the "predicted" values at older ages for the youngest birth cohort are extrapolations. No member of that cohort was older than age 45 during the survey period.

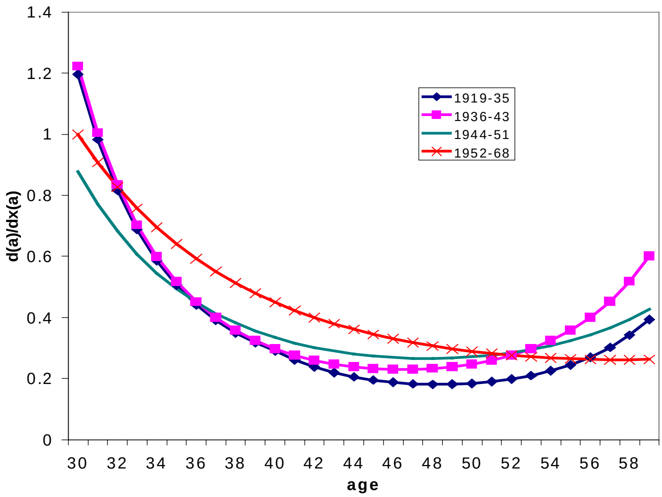


***Figure 4:*** *Ratio of general dependency ratio to extended household dependency ratio, four birth cohorts by age*

## 8. Specification 3: Substitution of survival for age effects

If forecasting or projecting is the ultimate objective of modeling living arrangements, then changing age effects in the proportion nuclear equation is an unwelcome complexity. One possibility is to project the age coefficients either by explicitly modeling the underlying processes or by some simple extrapolation. An alternative that is explored here is to identify one or more variables that captures the effects of aging, but are not subject to the shifts found with using years of age. An obvious candidate for estimating the proportion nuclear is the age-specific survival rate. The survival rate is closely related to many of the age-related phenomena that influence living arrangements. Most directly it bears on the probability of losing a spouse, the health status of seniors, and the longevity risks that affect their economic status. (Those who survive the high-risk ages towards the end of life are most likely to outlive their financial resources.)

Substituting the survival rate at each age for age, we regress the natural logarithm of the proportion nuclear on cohort dummies, the sex ratio at age 60, and the natural logarithm of the survival rate. Four broad cohort dummy variables are included, as above, to capture the effect of birth cohort. The survival rates are interacted with the birth cohort dummies to assess whether or not survival effects are changing across birth cohorts. In regression 1 we do not include the sex ratio of the birth cohort at age 60 to capture the surplus of immigrant males. In regression 2 the sex ratio is included.

***Table 3:*** *Estimates of cohort and survival effects for the proportion nuclear (60 and older)*

| Variable & Birth cohort | Regression 1 | | Regression 2 | |
|---|---|---|---|---|
| | Coefficient | Standard error | Coefficient | Standard error |
| Cohorts | | | | |
| 1893-1908 | -1.145 | 0.062 | -1.952 | 0.100 |
| 1909-1915 | -0.762 | 0.039 | -1.691 | 0.090 |
| 1916-1922 | -0.430 | 0.027 | -1.431 | 0.098 |
| 1923-1938 | -0.506 | 0.025 | -1.325 | 0.080 |
| Survival rate | | | | |
| 1893-1908 | 6.617 | 0.592 | 6.270 | 0.583 |
| 1909-1915 | 6.255 | 0.604 | 5.874 | 0.532 |
| 1916-1922 | 8.821 | 0.799 | 8.414 | 0.796 |
| 1923-1938 | 6.325 | 1.009 | 14.121 | 1.511 |
| Sex ratio | - | - | 0.726 | 0.068 |
| $R^2$ | 0.97 | | | 0.96 |
| N | 546 | | | 546 |

The results, presented in Table 3, are somewhat mixed. In regression 1, the impact of the survival rate is much more stable across birth cohorts than the effects of age presented above. With the exception of the third birth cohort, there is no discernible difference in the effect of survival rates on the proportion nuclear. The estimated cohort effect is broadly similar to the estimates reported above – a non-monotonic pattern persists with the proportion nuclear declining from the third to the fourth birth cohort.

The introduction of the sex ratio into the specification (regression 2) has a large effect on the model. The sex ratio has a significant positive effect as expected (Note 14). The cohort effect is monotonically increasing once the sex ratio is controlled. The estimated survival effects are sensitive to the inclusion of the sex ratio. The survival effect is much greater for the fourth birth cohort when the sex ratio is controlled than when it is not. In this formulation, then, the goal of finding a stable "age effect" remains elusive.

## 9. Projection model and projections

In order to project the proportion of seniors living in extended households we employ a simpler specification that assumes a constant trend in the cohort effect and a survival effect that is independent of birth cohort. Moreover, we use the log-odds of the proportion nuclear as the dependent variable so as to constrain the projected values of the proportion nuclear to between 0 and 1. The estimated results are:

$$\begin{aligned} \ln(n/(1-n)) = &-56.13 + 0.028 BYear + 8.16 \ln s + 1.46 SexRatio \\ &(3.77) \quad (0.002) \qquad (0.58) \qquad (0.076) \\ &N{=}546 \quad \bar{R}^2 = 0.86 \end{aligned} \tag{10}$$

The value of *ddx* is projected using the estimates for the 1952-68 birth cohort reported in Table 2. The proportion of adults aged 30-59 living in extended households is projected as the product of the projected values *ddx* and the proportion seniors living in extended households, as shown by equation (2). Projections of the age-distribution of Taiwan's population and its survival rates are taken from Lee et al. (2000) (Note 15).

The projections of the proportion living in extended households are presented in Figure 5. The age-specific proportions living in extended households are projected to drop fairly significantly at all ages except where the residual effect of Taiwan's unusual sex ratio persists. For sixty-year-olds, the drop is by about twenty percentage points between 2000-05 and 2030-35. The changes are smaller at older ages – about ten percentage points among those 85 and older.

For non-senior adults, the greatest decline in the proportion extended is also among the youngest adults. We project a drop of about 25 percentage points for 30 years olds and very little change for those in their late forties and fifties.

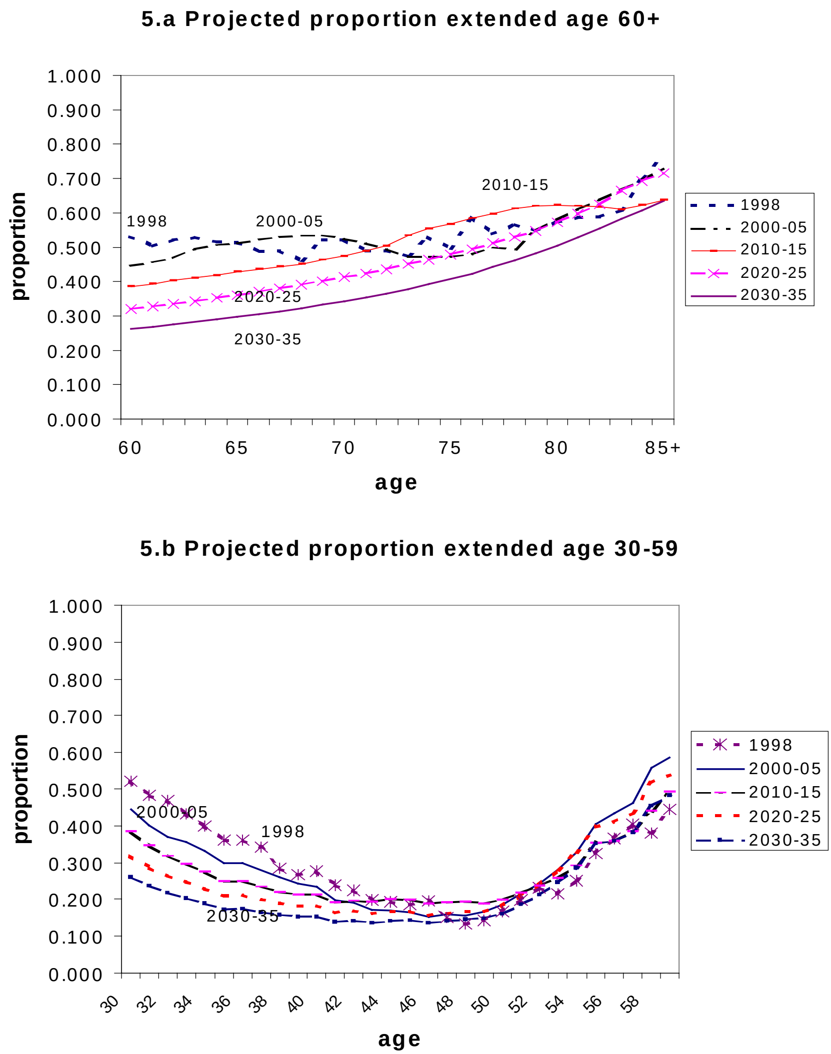


***Figure 5:*** *Projected proportion extended by age, 2000-2035*

The overall projection of extended living arrangements for seniors and for non-senior adults is shown in Figure 6. For seniors, we project a gradual decline in the proportion living in extended households. The proportion does not drop below 0.5 until 2010-15. By 2025-2030 the proportion is just below 0.4. The decline represents a departure from the relative stability that has persisted over the last twenty years. The projection of the proportion of non-senior adults living in extended families reverses the trend over the last twenty years. Between 1978 and 1998 the proportion 30-59 living in extended households has increased fairly steadily from a little less than 18 percent to almost 30 percent. Our projection is that the proportion extended will begin to decline again and drop to close to 20 percent by 2035.

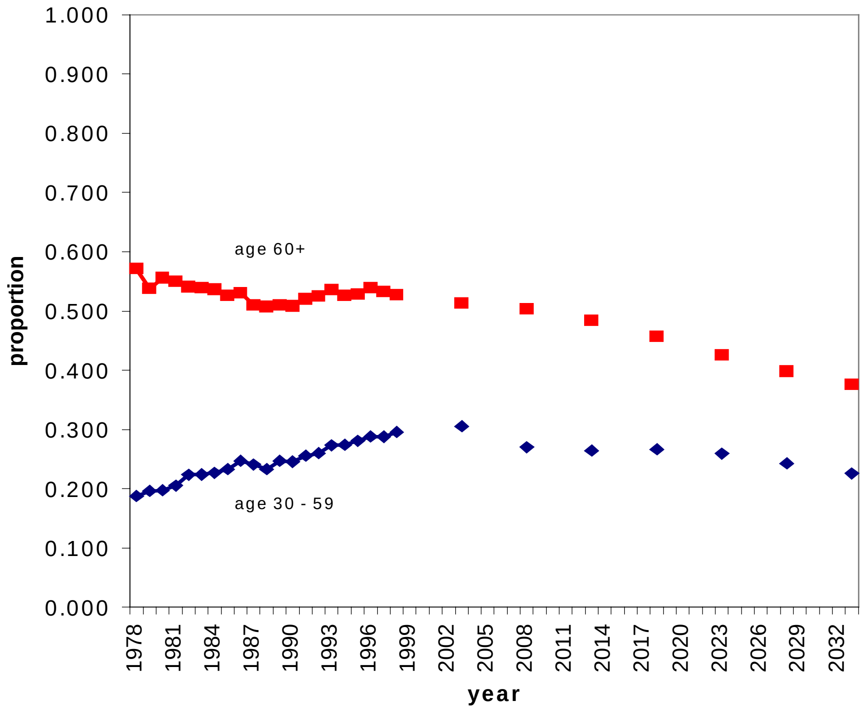


***Figure 6:*** *Projected proportion extended, persons aged 30-59 and 60+*

Why does the model project such a clear departure from the historical trend in living arrangements in Taiwan? Among seniors the large but temporary rise in the sex ratio led to a temporary elevation of the proportion living in nuclear households. As the sex ratio returned to more normal levels, the proportion living in nuclear households declined. In the future, changes in the sex ratio will not play an important role and the trend in the proportion of seniors living in extended households will be governed by the cohort and age effects, captured by the year of birth and survival variables in equation (10). Both the cohort and age effects are leading to a decline in the proportion of seniors living in extended households.

The trend in the 60+ age group is also affected by compositional changes. Over time the population is becoming increasingly concentrated at higher ages. Because the proportion living in extended households rises with age, the effect of age composition is to raise the proportion of seniors living in extended households. The composition effect is not sufficiently large, however, to outweigh the decline in the age-specific proportions of seniors living in extended households.

The trend in the non-senior age group is driven by the changes in the proportion of seniors living in extended households and changes in the relative dependency structure (*ddx*). During the period analyzed changes in age composition play a relatively minor role. Between 1978 and 1998 the upward trend in the proportion living in extended households was primarily among non-seniors in their early 30s and their 50s. Among those aged 35-49 little change occurred. Among the young non-seniors the rise in proportion living in extended households was tied to a rise in the proportion of seniors in their early 60s living in extended households. The relative dependency structure changed very little for the young non-seniors. Among older non-seniors, those in their 50s, the increase in the proportion in extended households was driven by a significant increase in *ddx*. The proportion of older seniors living in extended households declined significantly during this period, but changes in *ddx* more than offset the effect of the decline in extended families among seniors.

In the projection the decline in the proportion of non-seniors living in extended households is driven by the decline in the proportion of seniors living in extended households. The relative dependency structure is held constant in the projection and, thus, plays only a passive role in determining the proportion of non-seniors living in extended households. On the basis of the evidence at hand, we cannot dismiss the possibility that *ddx* will increase further because of additional decline in the net reproduction rate or further improvements in the survival rates among adults. If this does occur, the proportion of non-seniors living in extended households could continue to rise or decline at a slower pace than projected.

## 10. Conclusions

Population aging changes the availability of kin with whom intergenerational exchange of any form can take place. Co-residence is one form of this exchange that is influenced by age structure. As the number of seniors increases relative to the number of non-senior adults, at least one of the three outcomes must occur: (1) the number of seniors living in extended households must decline; and/or (2) the old-age dependency ratio within extended households must rise more rapidly than the old-age dependency ratio for the general population; and/or (3) the proportion of non-senior adults living in extended households must rise.

In Taiwan, no single outcome has dominated. Once we control for the effects of immigration on the availability of men, the proportion of seniors living in extended households is clearly in decline. The downward trend among young seniors, the cohort effect, appears to be relatively modest. Changes in the effects of aging appear to be more important. The proportion of seniors living in extended households rises significantly as seniors age. However, the effects of individual aging have moderated over time. Our purpose in this paper is not to explain why such changes have occurred and there are several obvious competing possible explanations. Seniors have higher income, they may be healthier, and the non-family support systems may have strengthened.

Because of the family linkages, one would expect the proportion of seniors living in extended households and the proportion of non-seniors living in extended households to move together. This occurred to some extent in Taiwan – among young seniors and young non-seniors. Among older seniors and non-seniors, however, the proportion living in extended households moved in opposite directions. While the proportion of older seniors living in extended households declined the proportion of older non-seniors living in extended household increased. The reason this occurred is because the dependency ratio in the general population rose relative to the dependency ratio in extended households. The number of older seniors increased relative to the number of older non-seniors, producing a rise in the proportion of non-seniors living with senior parents.

Despite the presence of factors driving a shift toward nuclear households, Taiwan actually experienced an increase in the proportion of adults living in extended households between 1978 and 1998. In part this occurred because of changes in the age composition of the population, in part because of a decline in the sex ratio to a more normal level, and in part due to the effects of aging on the proportion of non-seniors living in extended households.

Projections based on the model that we estimate here indicate that the forces favoring nuclear households may begin to dominate with the proportion of seniors and non-seniors declining gradually. This is a tentative conclusion, however, because considerable uncertainty remains about the relative changes in the age structure of the population and the age structure of extended households. Further increase in the proportion of non-seniors living in extended households would not be a surprising development.

A striking feature of the analysis is the resilience of extended living arrangements in Taiwan. The twenty-one year period spanned by our analysis, 1978 to 1998, was one of enormous social, economic, and demographic change in Taiwan. Those who reached their 60s in 1998 were much richer, more educated, more urbanized, and had substantially fewer surviving children than those who reached their 60s in 1978. Those who reached 60 in 1998 were somewhat less likely to live with their adult children than those who reached 60 in 1978, but the cohort trend is quite modest. At the same time, adult children were more likely to live with their senior parents than in the past. Although our projection model anticipates that the extended family will become less prevalent in the future, its decline is far from precipitous.

## Notes

1. Not all risks are downside risks. Investment risk may lead to an increase as well as to a decline in financial resources in any period. Even here, however, if elderly face repeated shocks the proportion whose wealth drops below any given level (say the level required for independent living) will increase over time (with age).

2. Annuities are subject to adverse selection, i.e., only those who expect to live a long life purchase annuities. This drives up the price to levels that are unattractive to individuals who do not expect to live to an unusually old age.

3. In the application of the model we use an open-ended age interval for those who are 90 and above.

4. A variety of factors will influence the errors introduced by this simplifying assumption. The surviving number of offspring to a cohort aged x is equal to the population aged x-g in a stationary population if the age distribution of offspring is symmetric. The greatest errors would arise when the population has been subject to large shocks. For example, given a sudden drop in fertility, the population aged *a-g* would be an underestimate of the number of surviving children of the high fertility cohort and an overestimate of the surviving children of the low fertility cohort.

5. There may be time effects as well. For example, fluctuations in the unemployment rate may influence the extent of co-residence for all birth cohorts and age groups. The analysis here is concerned with long run trends, however, and the effects of annual fluctuations are absorbed in the error term in the statistical analysis presented below.

6. Recall that the model is for a one-sex population. In a two-sex population, the laterally extended household would be comparable to all sons and their spouses living with both surviving parents and the stem family example would be comparable to a son and his spouse living with both surviving parents. The effect of mortality among senior married couples living in extended households is discussed below.

7. In our definition of extended families below, we include individuals living with their adult grandchildren irrespective of whether or not the children are surviving. Thus, to live in an extended family requires only that seniors had at least one child survive long enough to produce children.

8. See Cho and Yada 1994; Thornton and Lin 1994; and Mason et al. 1998.

9. The generation length of 30 is based on the average differences in age found in extended households between prime-age adult members and their co-resident parents. Young non-seniors are 30 years younger than their parents for all birth cohorts. The generation length is shorter for older non-seniors. By age 50 the generation length was about 21 years in 1978 and about 26 years in 1992 and 1998.

10. To calculate the value of *ddx* we extended the upper age group to 89+. *d(59)* is calculated as *N(89+)/N(59)* and *dx(59)* is calculated in a similar manner.

11. The coefficient for age 60 is constrained to equal 0.

12. Zimmer et al. (2002) present evidence that disability among the elderly, as measured by difficulty in walking and climbing, increased in Taiwan between 1993 and 1999. Whether this is true prior to 1993 is unknown.

13. The value for coefficient of the linear term for the 1936-43 cohort is -.203+(-.034)=-.237. The coefficient for the squared term can be calculated in similar fashion. The coefficients for the second birth cohort are not statistically different than the coefficients for the oldest birth cohort.

14. A more appropriate general formulation might measure absolute deviations in the sex ratio from 1 as either a large surplus of men or women would affect the proportion nuclear. In this particular instance, however, the only large deviations are positive.

15. Thanks to Tim Miller for providing the projections. Details of the projection are available in Lee et al. 2000. The projections assume that Taiwan will experience no immigration.

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