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Aging, Family Transfers, and Income Inequality

by

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Several recent studies (Deaton and Paxson 1997; Schultz 1997) conclude that population aging leads to an increase in consumption or income inequality. The results are based on a conventional model that analyzes compositional effects given inter-age differences in the means and variances of income (Lam 1997; Lam and Levison 1992). Depending on the response of public and private transfer systems, however, analysis limited to compositional effects may be misleading. The prospect of population aging has motivated the reform of many public transfer systems in Latin America, Europe, and the US. Although the ultimate effect of reform on their redistributive features remains uncertain, a recent analysis of OECD data led Gruber and Wise (Gruber and Wise 2001) to conclude that aging has led to a decline in the share of resources going to the elderly.[2] Similarly Razin et al. (2002) conclude that a rise in the overall dependency ratio is leading to a decline in social transfers. In contrast, Preston (1984) contends that elderly in the US were able to claim a disproportionate share of public resources as their numbers and political power increased.[3] The link to income inequality is completed by von Weizsacker (1995) who shows that depending on the response of public transfer aging can lead either to a rise or a decline in income inequality.

In many countries the familial support system is as important or more important than the public system. Familial support systems are similar to public support systems in that they are vulnerable to the same demographically induced "fiscal pressures." As the number of surviving pensioners increases and the number of workers declines within any family, either the workers must increase the share of their resources devoted to supporting their parents (higher "taxes") or parents must experience a decline in the share of their resources derived from relying on their children (reduced benefits). On theoretical grounds, however, the effect of aging on family transfers is uncertain. Given a sufficiently high degree of altruism, families may adjust their intergenerational transfers to distribute resources equally among all members. But if altruism is weak or intergenerational transfers are motivated by non-altruistic concerns, income differences

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[2] Gruber and Wise 2001 estimate that in OECD countries an increase in the share of the elderly population by 1 percent leads a rise in spending on the elderly by 0.47 percent.

[3] For an alternative interpretation of the U.S. experience see Becker and Murphy (1988).

within the family may widen in response to population aging (Altonji, Hayashi and Kotlikoff 2000; Becker and Tomes 1976; Cox 1987; Ermisch 2003; Frankenberg, Lillard and Willis 2002; Lillard and Willis 1997; McGarry and Schoeni 1997).

An important distinction between public and familial support systems is that the financial burden of a public system is spread across all workers or taxpayers, whereas the financial burden of a family system is concentrated only on those with surviving parents. As mortality conditions improve, providing familial support to the elderly becomes a responsibility shared by a larger share of prime age adults. Population aging leads to a 'broadening of the tax base' because the proportion of workers with a surviving parent increases. The effect of increased life expectancy is reinforced by the compression of the age of death. As the variance in the age of death declines so does the variance in the proportion of prime age adults with surviving parents.

Given these considerations it is far from clear how the familial support system and income inequality will respond to population aging. These issues are addressed in this paper by: (1) extending the standard model of income inequality to incorporate the responses in one aspect of the familial support system – forming multi-generational households; (2) developing a new empirical strategy for estimating the effects of aging, income, and other socio-economic variables on living arrangements; and, (3) applying the models to Taiwan, where public support systems are relatively under-developed but familial support systems are important.

Several studies have examined aspects of income inequality related to the ones addressed here. Schultz (1982) devised a decomposition method for analyzing the interaction between income, the number of adults per households, and the number of surviving children. Several studies have considered the effects on income inequality of the pooling of income by husbands and wives (Lam 1997; Lam and Levison 1992; Lehrer and Nerlove 1981; Liu and Chang 1987; Ogawa and Bauer 1996; Pong 1991). The effect of aging on income inequality incorporating the response of public transfer systems has been explored by von Weizsacker (1996). This paper complements these previous studies by incorporating the response of family support systems to changes in the age distribution of the population.

Although the major emphasis of this paper is methodological, the analysis of inequality in Taiwan is of interest in its own right. Taiwan has experienced rapid demographic and economic change. Life expectancy at birth increased from roughly 25 in 1900 to 78 for females and 72 for males in 1999. The total fertility rate declined from over 6 births per woman in the 1950s to replacement level by 1984. Since 1960 Taiwan's economy has been one of the fastest growing in the world. Yet despite these and other equally dramatic social changes, income inequality has been steady and low since the mid-1960s. Between 1953 and 1964 the Gini coefficient dropped from about 0.55 to 0.33, giving it one of the lowest levels of income inequality in the developing world (Liu 1982). Between 1978 and 1998, the period analyzed in this paper, income inequality declined by about 10 percent.[4] How Taiwan achieved such rapid economic growth but avoided rising income inequality is an important development policy issue (World Bank 1993). The answer is to be found, in part, in the responses of the familial support system.

The paper is organized as follows. Section I briefly describes a new model that is useful for analyzing aggregate variables, e.g., income inequality. Section II extends the

[4] Based on the coefficient of variation of income per adult.

standard method for analyzing the compositional effect of age structure on income inequality to incorporate the effects of income pooling that occur when family members establish multi-generational households. Section III models the effects of age structure, income, and other variables on living arrangements. Section IV uses data from Taiwan to estimate the effects of demographic and economic variables on living arrangements and income inequality. Section V discusses reservations and problems and Section VI concludes.

I. The Overlapping Families (OLF) Model

Aggregate economic/population models currently come in two forms: overlapping generation (OLG) models and demographic models. Although OLG models can consist of many generations, two or three generation models are frequently used. The three-generation OLG model illustrated in Figure 1 captures both age structure and intergenerational family structure. Its simplicity is a distinct advantage in theoretical work, but the crude fashion in which age structure and intergenerational family structure are captured is a disadvantage in empirical work.

The demographic model, represented by a standard age pyramid, uses detailed information about age structure. This offers advantages in empirical work or in simulating dynamic effects of population change. The disadvantage of the demographic model is that the intergenerational family structure is lost.

Figure 1. Two Perspectives on Population: Overlapping Generations Model and Demographic Model

The overlapping families (OLF) model, introduced here, is a synthesis of these two approaches capturing the advantages of both: detailed age structure and intergenerational family structure. Family structure in the OLF model is based on an unrealistic, but useful assumption – that all births are to women (or persons in a one sex version) of age g. Persons of age a, thus, have children aged a-g and parents aged $a+g$.

The population is readily sub-divided into generations. The youngest generation consists of those who have not yet given birth, i.e., those younger than age g. The next youngest generation consists of those who have given birth but whose children have not. They are between the ages of g and $2g$. Older generations are defined in similar fashion. Depending on life expectancy and the generation length the number of generations usefully distinguished could be more than 3, but in the population we analyze distinguishing three generations appears to be sufficient. In Figure 2 the generation length is 30 years and we use the titles *children, workers,* and *pensioners* to distinguish the three generations represented. We should emphasize that these titles are used to capture the main features of the life cycle, but they do not embody assumptions about labor force, pensions, or other economic or social behavior.

Figure 2. Overlapping Families Model

The population can also be sub-divided into mutually exclusive, exhaustive family cohorts. A family cohort consists of all workers aged a, children aged a-g, and pensioners aged $a+g$. We index the family cohort using the age of the members who are workers. Two family cohorts are shown in Figure 2: the family cohort that consists of workers aged 30-34, pensioners aged 60-64, and children aged 0-4 and the family cohort consisting of workers aged 45-49, pensioners 75-79, and children 15-19. Although the model is represented using five-year age groups, any age grouping can be used in principle. In this application we employ single-year age groups.

The OLF model used here is restricted to two generations: workers and pensioners. The analysis does not consider the effects of children on income inequality, although this would be a useful extension. How to do so is a controversial subject, however, because children are a choice variable. Schultz (1997) suggests that inequality in income per adult, the measure employed here, is preferred so as to avoid endogeneity problems.

II. Income Inequality

Income inequality is measured as the coefficient of variation (*CV)*, which has well-known properties and has been used previously to analyze the effects on income inequality of household composition (Lam 1997) and aging (von Weizsacker 1995). The analysis uses the OLF model and proceeds in two steps. First, we consider income inequality for family cohorts. Second, results for family cohorts are used to construct estimates for the population as a whole.

The second step in the analysis involves no innovation. Given the mean ($\bar{Y}_a$) and variance ($V_a(Y)$) of income for family cohorts, income inequality for the population depends only on compositional effects. Applying the conventional formula to family cohort data:

$$\begin{aligned} V(Y) &= \sum_{a=g}^{2g} u_a V_a(Y) + \sum_{a=g}^{2g} u_a (\bar{Y}_a - \bar{Y})^2 \\ \bar{Y} &= \sum_{a=g}^{2g} u_a \bar{Y}_a \\ CV &= \sqrt{V} / \bar{Y} \end{aligned} \tag{1}$$

where *V(Y)* is the variance in per adult household income for the population, $\bar{Y}$ is the mean income per adult, u_a is the proportion of the adult population belonging to the family cohort of age *a*, and CV is the coefficient of variation. The values are summed over the age-interval for the worker generation because, as noted above, the ages of the worker generation are used to index family cohorts.

Before equation (1) can be employed, however, we must determine how the family cohort variables $V_a(Y)$ and $\bar{Y}_a$ are influenced by the age composition and living arrangements of family cohorts. Each family cohort consists of two generations, workers aged *a* and pensioners age *a+g* with mean incomes of $\bar{Y}_a^k$ and $\bar{Y}_a^p$, respectively, and variances of $V_a(Y^k)$ and $V_a(Y^p)$, respectively. These values are taken as given throughout the analysis. The proportion of family members who are workers is designated by m_a^k and the proportion who are pensioners by m_a^p where $m_a^k + m_a^p = 1$.

The mean income of the family cohort, given by:

$$\bar{Y}_a = m_a^k \bar{Y}_a^k + m_a^p \bar{Y}_a^p \tag{2}$$

is a simple weighted average of the mean incomes of the worker and parent generations. It is independent of living arrangements.

The variance of income for the family cohort is influenced by both age structure and living arrangements. It is this relationship that is key to understanding how changing living arrangements influence income inequality within the family cohort and the population. The remainder of this section analyzes how changes in age structure and living arrangements influence $V_a(Y)$.

Analysis is greatly eased through a number of simplifying assumptions that abstract from details that are less important to the issues at hand. The implications of relaxing these assumptions are discussed in section VI. First, we assume that households come in one of two forms. Nuclear households consist of a single worker or a single pensioner. Extended households consist of a single worker and a fractional portion of a pensioner. Second, the decision to establish an extended household is independent of the income of either the worker or the pensioner. The incomes of workers and pensioners who belong to the same family are assumed to be correlated. A positive correlation, for example, leads to a positive covariance between the incomes of workers and pensioners living in extended households, but this does not arise because income influences the decision to co-reside.

Given these simplifying assumptions, the variance of income per adult is given by:

$$\begin{aligned}
&V_a(Y) = w_1 V(Y_a^k) + w_2 V(Y_a^p) + w_3 C(Y_a^k Y_a^p) + w_4 (\bar{Y}_a^k - \bar{Y}_a^p)^2, \text{ where:}\\
&\quad w_1 = m^k - m_x^k m_x^p m_x \geq 0\\
&\quad w_2 = m^p - m_x^k m_x^p m_x \geq 0\\
&\quad w_3 = 2 m_x^k m_x^p m_x \geq 0\\
&\quad w_4 = m^k m^p - m_x^k m_x^p m_x \geq 0.
\end{aligned} \tag{3}$$

The coefficients, w_i, and the variables that determine them all vary across family cohort but the *a* has been dropped to ease notation. The coefficients depend on the age-structure of the family cohort – the proportion of members who are workers (m^k) and pensioners (m^p); the age-structure of extended households – the proportion of members who are workers (m_x^k) and pensioners (m_x^p); and, the proportion of cohort family members who live in extended households (m_x). The four coefficients are non-negative. The first three coefficients sum to one. Hence, the variance in per adult income is determined, in part, as a weighted average of the variances of the incomes of the family members and the co-variance between their incomes ($C(Y_a^k Y_a^p)$). This is called the difference-in-variances component. The fourth right-hand-side term captures the effect of differences in the average incomes of workers and parents, called the difference-in-means component.

The effect of changes in the age structure of the family cohort is identical to that found in conventional models (Lam 1997). An increase in the relative size of the age group with the higher income variance leads to an increase in the difference-in-variance component. A shift towards a more balance age distribution leads to an increase in the difference-in-means component. Formally, the partial effect of aging as measured by an increase in the proportion of family members belonging to the pensioner generation is given by:

$$\partial V_a(Y)/\partial m_a^p = V_a(Y^p) - V_a(Y^k) + (1-2m_a^p)(\bar{Y}_a^k - \bar{Y}_a^p)^2 \tag{4}$$

The effect of aging on $V_a(Y)$ given living arrangements is illustrated by Figure 3, constructed using average values for Taiwan during the 1978 to 1998 interval. The 'nuclear household line' shows how $V_a(Y)$ would vary with age structure for the 'typical' family cohort were there no intergenerational pooling of income. The relationship follows an inverted U with the peak occurring when about one-quarter of all adults belong to the pensioner generation.

Figure 3 about here. Aging and Income Inequality, Family Cohort

The lower line in Figure 3 shows the effects of age composition on income inequality incorporating the average observed patterns of intergenerational co-residence for Taiwan. Co-residence reduced income inequality by about 18%. Note that the pooling effect and the age composition effect are additive. The age composition is identical whether or not families are living in extended households. Likewise, the effect of pooling is to reduce income inequality by about 18% irrespective of the family cohort's age composition.

The effect of living arrangements is captured by the multiplicative term $m_x^k m_x^p m_x$ in equation (3), which we call the pooling effect. The pooling effect increases with the age balance in extended household membership ($m_x^k m_x^p$), reaching a maximum when there are equal numbers of workers and pensioners, and with the proportion of family cohort members who live in extended households (m_x).

Defining $R = V(Y^p)/V(Y^k)$ and r as the correlation between the income of parents and workers, equation (3) can be re-written as:

$$V(Y) = (w_1 + w_2 R + r\sqrt{R}w_3)V(Y^k) + w_4(\bar{Y}^k - \bar{Y}^p)^2 \tag{5}$$

The partial effect of an increase in the proportion living in extended households is:

$$\partial V(Y)/\partial m_x = m_x^k m_x^p (2r\sqrt{R} - 1 - R)V(Y^k) - m_x^k m_x^p (\bar{Y}^k - \bar{Y}^p)^2 \leq 0. \tag{6}$$

If $r = R = 1$, the term $2r\sqrt{R} - 1 - R$ is zero; otherwise it is negative. Hence, the partial effect of an increase in the proportion of the family cohort living in extended households is to reduce the variance in income.

The partial effect of an increase in the proportion of pensioners in extended households is given by:

$$\partial V(Y)/\partial m_x^p = (1-2m_x^p)m_x\left(2r\sqrt{R}-1-R\right)V(Y^k)-(1-2m_x^p)m_x(\bar{Y}^k-\bar{Y}^p)^2. \qquad (7)$$

The partial effect is ambiguous. If the share of pensioners in extended households is sufficiently high, further increases in their share will lead to a decline in the variance of income. If the share of pensioners is less than one-half, however, the partial effect is unambiguously negative. As just noted $2r\sqrt{R}-1-R \leq 0$. Hence, $\partial V(Y)/\partial m_x^p < 0$ for $m_x^p < 0.5$. As an empirical matter, the partial effect is negative for about 90% of the observations in Taiwan.

With the effects of aging and co-residence on income inequality established, the following section can address the inter-dependence between population aging and the co-residence variables. The following section analyzes how the proportion of family members living in extended households and how the age-structure of extended households is influenced by changes in the age-structure of the family cohort.

III. Co-residence

Our model of co-residence draws on the extensive literature on living arrangements and intergenerational transfers, but the effects of aging are emphasized. Co-residence provides an efficient means for carrying out inter-generational transactions (Ben-Porath 1980). The transfer of time is facilitated in obvious ways by co-residence. To the extent that shirking, moral hazard, or adverse selection are problems in family exchange, co-residence may facilitate monitoring and increased efficiency in the allocation of family resources of any kind. Co-residence also allows families to realize gains from scale economies in home production and the consumption of household public goods (Ermisch 2003).

These gains are achieved at a cost. Co-residence restricts the geographic mobility of family members. By co-residing family members sacrifice their privacy and perhaps control over personal resources. Thus, family living arrangements may change over time as the importance of intergenerational transfers change; as the consumption of goods and services produced within the household shift relative to goods and services that are purchased in the market place and consumed by individuals; or as the interests of generations converge or diverge.

Living arrangements are not entirely a reflection of choice, however. Observed patterns may reflect kinship availability (Ruggles 1987; Wachter, Hammel and Laslett 1978). Only multi-generation families can establish multi-generation households. Moreover, the age structure of multi-generation household may be influenced by the age structure of multi-generation families.

There is an extensive empirical and theoretical literature on multi-generation living arrangements (Bachrach 1980; Chevan and Korson 1975; Kobrin 1976;

Macunovich et al. 1995; Mason and Lee 2003; Michael, Fuchs and Scott 1980; Soldo 1981; Wister and Burch 1983; Wolf 1995) and on intergenerational transfers (Altonji et al. 2000; Becker and Tomes 1976; Costa 1997, 1998; Cox 1987; Ermisch 2003; Frankenberg et al. 2002; Lillard and Willis 1997; Martin 1989; McGarry and Schoeni 1997; Palloni 2001; Thornton and Lin 1994). Some of the issues addressed in the literature are relevant to Taiwan's experience while others are not. Taiwan provides a poor "laboratory" for assessing the impact of public transfer systems on family transfers, because social support systems for the elderly were relatively under-developed in Taiwan throughout the period of analysis. On the other hand, Taiwan experienced very substantial changes in income and demographic variables making it an ideal place to study their influence on living arrangements.

The relationship between income and intergenerational transfers in competing theoretical models is succinctly summarized by Lillard and Willis (1997), whereas Palloni (2001) provides a recent review of the literature on income and living arrangements. Empirical studies of the West find a relatively weak relationship between income and living arrangements (Borsch-Supan et al. 1992; Michael et al. 1980; Schwartz, Danziger and Smolensky 1984) and a weak relationship between income and transfers (Altonji et al. 2000). Income appears to have a much stronger impact on transfers in developing Asia than in the West (Frankenberg et al. 2002; Lillard and Willis 1997). Income also had a much stronger effect on living arrangements during the first few decades of the 20th Century in the U.S. Costa (1997; 1998) concludes that higher income was primarily responsible for the decline in the extended U.S. family.

There is an extensive literature on the effect of kin availability on living arrangements. Historical studies of the West have addressed whether the low prevalence of extended households in the past was a consequence of the effects of high mortality on kinship availability (Ruggles 1987; Wachter et al. 1978). Many studies have found that the probability that older adults lived in extended households increases with the number of surviving children (Palloni 2001).

These issues are of particular interest here because of the obvious connections between aging and kinship availability. We use an analytic approach to show how population aging influences the kinship group, i.e., the *family*. The model used here is similar to but much simpler than other models used to study kinship availability (Freedman et al. 1991; Lin 1994; Ruggles 1987; Wachter et al. 1978; Wolf 1995).

Two aspects of the family cohort are relevant to income inequality: (1) the proportion who are members of multi-generation families (m_x^F); and, (2) the share of pensioners in multi-generation families. Here and in the remainder of this section, the family cohort index *a* is dropped to simplify notation, but all variables and parameters are specific to a family cohort. It is more convenient to analyze the old-age dependency ratio of multi-generation families (D_x^F), which is a monotonic transform of the share of pensioners.

In the following sections we show, first, that under many, but not all, circumstances population aging leads to an increase in the proportion of the family cohort belonging to multi-generation families. Second, under many, but not all, circumstances population aging leads to an increase in the old-age dependency ratio within multi-generation families. Third, the effects depend on whether fertility or mortality change underlie population aging.

Proportion in multi-generation families

For any family cohort closed to migration[5] the population of the pensioner generation is given by *M(a+g,t)=s(a,t)M(g,t-a),* where *s(a,t)* is the survival rate. The population of the worker generation is *M(a,t)=f(a,t)N(g,t-a),* where *f(a,t)* is similar to the net reproduction rate. Dropping the age and year indexes to simplify notation, the age composition of the family cohort is measured by the dependency ratio, $D = s / f$. The proportion of persons belonging to a multi-generation family is given by:

$$m_x^F = \frac{a\, s + b\, f}{s + f} \tag{8}$$

where a is the proportion of pensioners who belong to a multi-generation family, i.e., the proportion with at least one surviving offspring, and b is the proportion of workers who belong to a multi-generation family, i.e., the proportion with at least one surviving parent.

As is apparent from equation (8), changes in the age distribution have a compositional effect that depends on whether pensioners or workers are more likely to have intergenerational links. The proportion with intergenerational links depends, in turn, on survival and fertility. Members of the pensioner generation will have no intergenerational links if they are celibate or pre-deceased by their children. Improvements in survival at the working ages will lead to a decline in the proportion of pensioners who are pre-deceased by their children ($\partial a / \partial s_k > 0$), where s_k is the proportion surviving from birth to age *a*. Reductions in fertility may lead to an increase in the proportion who are celibate ($\partial a / \partial f > 0$).

There is no obvious reason why the proportion of workers belonging to multi-generation families would be influenced by fertility. Changes in survival, however, have a direct influence on the proportion of workers with a surviving parent. In a one sex population the relationship between the survival rate *s* and the proportion of workers with a surviving parent is trivial, i.e., $b = s$. This systematically under-estimates the proportion of workers with a surviving pensioner in a two-sex population unless the mortality of mothers and fathers are perfectly correlated. For this reason we relax the one-sex assumption maintained elsewhere in this paper. It can be shown that if *s* is the survival rate and f is the correlation between the survival of mothers and fathers, then:[6]

$$\begin{aligned} b &= s + s(1-s)(1-f) \\ \frac{\partial b}{\partial s} &= 1 + (1-2s)(1-f) \geq 0 \text{ if } f \geq 0 \\ h_b &= \frac{1+(1-2s)(1-f)}{1+(1-s)(1-f)},\ 0 \leq h \leq 1 \text{ if } f \geq 0, \end{aligned} \tag{9}$$

where h_b is the elasticity of b with respect to *s.* If the survival of husbands and wives is positively correlated or independent, then an increase in the survival rate leads to an

[5] Immigration is discussed below.

[6] This assumes that the mortality of parents and offspring are independent. Proof available from authors by request.

increase in the proportion of workers with surviving parents. Likewise, the elasticity is positive but less than one.

The partial effect of aging on the proportion of the family cohort who are members of multi-generation families can now be assessed using equation (8). First, consider the effect of aging due to an increase in survival. Using a' to represent $(\partial a/\partial s_k)(\partial s_k/\partial s) > 0$ and b' to represent $\partial b/\partial s > 0$:

$$\frac{\partial m_x^F}{\partial s}\frac{\partial s}{\partial D} = \frac{(a-b)+a'(s+f)s+b'(s+f)}{(s+f)^2} f^2 > 0 \text{ if } a > b. \qquad (10)$$

Note that even if $a < b$ the partial effect will be positive if the effects of survival on a and b are sufficiently great.

Second, consider the effect of aging due to an increase in fertility. Letting a' represent $\partial a/\partial f$:

$$\frac{\partial m_x^F}{\partial f}\frac{\partial f}{\partial D} = \left[\frac{(a-b)f - a's(s+f)}{(s+f)^2}\right]\left[\frac{f^2}{s}\right]. \qquad (11)$$

If $a > b$ and a' is small, i.e., fertility has a negligible effect on rates of celibacy, aging due to fertility decline also leads to an increase in the proportion of the family cohort belonging to multi-generation families. However, if fertility decline leads to a sufficiently large increase in the proportion of the pensioner generation who are childless, the proportion of persons belonging to multi-generation families will decline.

The effect of aging turns to a substantial degree on whether pensioners or workers are more likely to belong to multi-generation families. In Taiwan, the subject of our empirical analysis, and three other East Asian populations recently studied by Hermalin and Myers (2002), the proportion of persons 60 and older with a surviving offspring ranged from 0.956 to 0.961 in the 1990s. In these countries, survival rates are high and rates of celibacy are low. In other contexts, much lower proportions have been observed. The effects of high mortality are illustrated by India's experience.[7] In a 1956 survey of rural India, 22 percent of women 65 and older reported that they had no living *sons (Collver 1963)*. The US experience illustrates the potential importance of fertility. Of US women aged 45-59 in 1900, 1925, 1950, and 1975, 14, 22, 20, and 9 percent, respectively, had never given birth (Pullum 1987).

Based on surveys conducted in Japan, Korea, and the US in the late 1980s and early 1990s, of currently married couples 30-59, 21 percent, 27 percent, and 17 percent of the wives did not have a surviving pensioner, respectively; 27 percent, 29 percent, and 21 percent of the husbands has no surviving pensioner (Rindfuss et al. forthcoming). Clearly in high mortality populations much higher values would be found.

Using the notation employed here, values of a range from 0.78 to 0.96 with the high value characteristic of East Asia. Values of b range from 0.71 to 0.83 with the lowest values found in East Asia. None of the estimates of b, however, are drawn from a high mortality setting where lower values would be found. These estimates of a and b are averages for prime-age adults and seniors, but the values may vary substantially with the age of the family cohort. The proportion of workers with a surviving pensioner will decline with the age of workers. The proportion of pensioners with a surviving

[7] Only 3% of women had never given birth (Collver 1963).

offspring will decline as well, but given the typical age pattern of mortality, the decline will be more rapid for workers than for pensioners. Thus, the difference between a and b should increase with the age of the family cohort.

In East Asia, it appears that population aging would lead to an increase in the proportion of most family cohorts belonging to multi-generation families thus increasing the availability of candidates for multi-generation extended households. There is some overlap, however, and the opposite may hold for some young family cohorts.

Age structure of multi-generation families

The second availability measure relevant to this analysis is the age structure of multi-generation families, designated by D_x^F. The relationship between the age structure of the population (D) and the age structure of multi-generation families (D_x^F) is given by:

$$D_x^F = \frac{a\,s}{b\,f} = \frac{a}{b}D. \tag{12}$$

The partial effect of a rise in the dependency ratio due to a change in the survival rate is given by:

$$\begin{aligned}\frac{\partial D_x^F}{\partial s}\frac{\partial s}{\partial D} &= \left[\frac{a}{b}\frac{1+h_a-h_b}{f}\right]f \\ &= \frac{a}{b}(1+h_a-h_b) \geq 0,\end{aligned} \tag{13}$$

where h_a is the elasticity of a with respect to *s*, which is greater than zero, and h_b is the elasticity of b with respect to *s*, which is between zero and one. In the polar case, h_a equal to zero and h_b equal to 1, fertility decline has no effect on the composition of multi-generation families. In more realistic cases, an increase in the dependency rate due to improving survival leads to a rise in the dependency ratio in multi-generation families.

Letting $a' = \partial a / \partial f$, the partial effect of a rise in the dependency ratio due to a rise in fertility is given by:

$$\frac{\partial D_x^F}{\partial f}\frac{\partial f}{\partial D} = \frac{a}{b} - \frac{a'}{b}f. \tag{14}$$

The partial effect is positive unless fertility decline is accompanied by a significant rise in childlessness.

Additional effects of survival and fertility

The effects of survival and fertility analyzed in the two preceding sections are limited to availability effects, i.e., changes in the demographic characteristics of the kinship group or the family. There are other aspects of survival and fertility beyond their influence on availability.

Changing survival rates are accompanied by changes in health status that will influence the extent to which members of the pensioner generation can maintain independent living arrangements. Whether or not longer life expectancy is leading to improvements in health status is a complex issue about which there is only limited information in Taiwan (Zimmer, Martin and Chang 2002). In addition, improvements in survival rates will also raise the average age at which individuals become widowed possibly influencing the preference for living with children.

Fertility influences the average size of the sibset to which workers belong. To the extent that multi-generation households consist of only one sibling and his or her pensioners, a common arrangement in East Asia, fertility decline will result in an increase the probability that a member of the worker generation is living in an extended household.

Immigration

The simple model presented in this section presumes a closed population. International migration may influence the prevalence and age composition of multi-generation families in several ways. First, either out-migration by residents or in-migration by foreign groups will lead to a decline in the proportion of the resident population who are members of inter-generational resident families unless the migration involves multi-generation families. Second, if immigration is highly age- and sex-specific, it may lead to an imbalance in sex ratio, interfere with the formation of marital unions, and lead to a rise in childlessness for members of the surplus sex. Taiwan experienced an unusual immigration shock that allows us to assess whether sex-biased immigration can influence living arrangements and, thereby, income inequality. *Circa* 1950 Taiwan's population increased by 15% or more by the arrival of over one million mainland Nationalists. The immigrants were predominantly men of military age so that there is a large surplus of males in many birth cohorts.

IV. Empirical Analysis

The analysis of Taiwan consists of two distinct elements. First, we estimate the effects of aging and other variables on living arrangements. Second, we use simulation techniques to assess the effect of aging on income inequality.

Description of the Data

The empirical analysis is based on synthetic panel data constructed using repeated cross-sectional household survey data for Taiwan collected annually from 1978 to 1998. The primary data set consists of mean values for single-year family cohorts for 21 consecutive years. The generation length g is set to 30 years. This estimate is based on the difference in the average age of different generations observed in extended households. The generation length is quite stable across time for family cohorts under the age of 50.[8] Thus, the data consists of 30 family cohorts for each year consisting of workers aged 30-59 and pensioners aged 60-89. Those aged 90 and older are included with the 89-year-old pensioners. Thus, the data set consists of 21 years times 30 age groups, yielding 630 observations. The oldest members of the worker generation were born in 1923 and the youngest in 1968, but for these two groups we have only one observation. Family cohorts with workers born near the middle of the span of birth years covered (1955) can be tracked over the entire 21-year span of the data. The oldest members of the pensioner generation were born in 1888 and the youngest were born in 1938.

[8] For older cohorts the age difference declines due to the effects of mortality on the survival of pensioners leading to an under-estimate of the generation length.

The mean values of most variables that compose the synthetic panel are constructed from the Survey of Family Income and Expenditure in Taiwan (FIES), also known as the Survey of Personal Income Distribution in Taiwan until 1993. The number of household 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 interviewed. These are not panel data, but repeated cross-sections.

Respondents are classified as members of nuclear or extended households in the following manner. All household members related to the head 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. A second generation consists of all those reporting parent, aunt, or uncle of the head. Any household that includes adult members (30 or older) belonging to two different generations is classified as an extended household. Marital status is not a factor in the classification scheme.

An attractive feature of the FIES is that household income is assigned to members of the household. Although there is a residual category for income that cannot be assigned to an individual, this category is rarely used. Consequently, we can calculate income characteristics separately for the worker and pensioner generations within extended households. Income is measured by total current income. The means and variances of income are all based on income per adult. All means and variances are weighted by the number of adults in the household or sub-unit.[9]

Nuclear households are designated as pensioner households or worker households based on the age of the primary income earner. If he or she is 60 or older the household is classified as pensioner. In extended households, the adult members are classified as pensioners or workers based on the generation to which they belong. The members of the youngest adult generation in the household are designated as workers; the members of all other adult generations are designated as pensioners.

The definitions of variables and descriptive statistics are provided in the appendix.

Statistical Analysis of Living Arrangements

Regression estimates for the proportion of family cohort members living in extended households are reported in Table 1 and for the dependency ratio in extended households in Table 2. All regressions are estimated using ordinary least squares. For each dependent variable results from four specifications are reported. Specifications 1 and 2 capture the effects of age structure using the dependency ratio (specification 1) and its square (specification 2). Specifications 3 and 4 introduce adult survival into the model in order to analyze the separate effects of fertility and survival.

In all specifications we include the year of birth of the worker generation of the family cohort (*Cohort),* the natural log of the average earnings of members of the worker generation (ln *labwork*), the ratio of the average earnings of the pensioner generation to the average earnings of the worker generation (*Incrat),* the sex ratio of the pensioner generation at age 60 (*Sexratio),* and single-year age dummy variables. The coefficients of

[9] For a discussion of issues that arise in measuring income inequality see Lam 1997 or Schultz 1997.

the age dummies are not reported in the table. Standard errors are given in parentheses. Coefficients set in bold face are statistically different from zero at the 1% level.

Table 1 about here.

The coefficient of log worker earnings captures the effect of general increases in earnings because the ratio of worker earnings to pensioner earnings is controlled. The statistically significant negative coefficient is consistent with the standard view that higher income leads to an increase in the demand for privacy. At higher incomes individuals are willing to give up the gains that arise from scale economies or public goods captured by extended households. A rise in the earnings of pensioners relative to workers has no discernible effect on the proportion of family members living in extended households. The estimated coefficient is small and statistically insignificant. Thus, we find no support for the altruism hypothesis. Note, however, that the economic status of the pensioner generation may be poorly measured by current earnings. The effects of pensioners wealth might be more supportive of the altruism hypothesis as has been the case in other studies (Borsch-Supan, McFadden and Schnabel 1996). Adequate wealth measures by generation are not available from the data.[10] The cohort effect is positive and the *Sexratio* has a statistically significant negative effect as anticipated. A lower proportion of cohorts with substantial surplus males were living in extended, multi-generation households.

Based on specifications 1 and 2, the dependency ratio has a positive, statistically significant effect on the proportion of the family cohort living in multi-generation households. The elasticity rises gradually with the dependency ratio. This provides support for a central feature of this paper – that aging may lead to a reduction in income inequality by encouraging greater co-residence. Specifications 3 and 4 address whether the source of aging, fertility or mortality decline, matters. The analysis presented above shows that the effects on availability of a decline in fertility or an increase in survival are not identical. Moreover, changes in fertility and survival may influence co-residence in ways other than through their effects on kinship availability.

The issue is addressed empirically by including both the dependency ratio and the proportion of the pensioner generation surviving. Given ln *s*, the coefficient(s) of the $\ln D$ terms give the partial effect of an increase in *D* due to fertility decline. A rise in the dependency ratio due to fertility decline produces an increase in the proportion of family cohort members residing in extended households. Again the elasticity rises modestly with the dependency ratio.

The estimated coefficients of the ln *s* terms in specifications 3 and 4 are partial effects controlling for the dependency ratio and the net reproduction rate because $\ln D = \ln s - \ln f$. Two distinct effects related to survival are captured by the coefficients. The first is that the availability effects of survival are different than the availability effect of fertility. The second is that changing survival may capture other changes correlated with survival but unrelated to availability. An example is that changes in health status may be captured by survival rates, although the evidence on how health

[10] Home ownership is a poor measure of wealth for our purposes in Taiwan because of the common practice of transferring ownership from pensioner to child prior to the death of the pensioner or at the death of the patriarch (Hermalin, Chang and Roan 2002).

status effects living arrangements is mixed (Borsch-Supan et al. 1992; Palloni 2001) as is the evidence regarding health trends in Taiwan (Zimmer et al. 2002). Thus, two possible interpretations of the coefficients are possible. One is that the availability effects of increased survival are much smaller than the availability effects of fertility decline. The second possibility is that direct effects are being captured. For example, improvements in health associated with increased survival are leading to a decline in extended living arrangements. In the absence of additional information further interpretation of the results is unwarranted.

Noting that $\ln D = \ln s - \ln f$, the elasticity of m_x with respect to s is equal to:

$$\begin{aligned}\frac{\partial \ln m_x}{\partial \ln s} &= b_1 + b_3 + 2(b_2 + b_4)\ln s - 2b_2 \ln f \\ &= 0.164 - 0.014\ln s - 0.078\ln f.\end{aligned} \tag{15}$$

where b_i is the *ith* estimated coefficient of specification 4 in Table 1. The elasticity declines modestly as the survival rate and the fertility rate increase, but it is positive for all values of *s* and *f* that we observe. The bottom line is that aging due either to fertility decline or survival increase leads to a rise in the proportion living in extended households, but aging due to fertility decline has a larger effect than aging due to a decline in survival.

Analysis of D_x reported in Table 2 uses the same specifications as used for m_x. The cohort effect is negative and statistically significant. Average earnings has a statistically significant positive effect on the proportion of seniors in extended households, whereas a rise in the earnings of pensioners relative to workers leads to a decline in the proportion of seniors in extended households. A large surplus in the male pensioner population leads to a decline in the proportion of pensioners in extended households.

Table 2 about here.

The elasticity of D_x with respect to *D* is close to one in all specifications. In the constant elasticity specifications (1 and 3), the elasticity is not significantly different than one. In the specifications that allow for a changing elasticity (2 and 4), the elasticity is close to one when evaluated at the mean, but increases gradually with the dependency ratio. This is consistent with a very simple demographic model of the age composition of extended households. An increase in the dependency ratio due to fertility decline leads to a somewhat smaller increase in the dependency ratio in extended households than does an increase in the dependency ratio due to an improvement in survival.

Counter-factual Analysis of the Historical Trends in Co-Residence and Income Inequality

The historical implications for living arrangements and income inequality of changes in income, aging, fertility, and mortality are assessed using simulation techniques. We construct a series of counter-factual simulations that control for selected determinants while allowing other variables to change in accordance with their observed historical trend. First we vary the values of one or more of the independent variables, e.g., income or fertility. Where appropriate we determine how age structure would have changed under the counter-factual simulation. Using the results reported in Table 1 we estimate how living arrangements would have evolved given the counter-factual values. The counter-factual variance and mean incomes for each family cohort are calculated. Finally equation (1) is used to obtain a counter-factual estimate of *CV*, our measure of income inequality. All calculations employ the observed values for the variance and mean of income of worker and pensioner generations by single years of age for the year in which the calculations are being made.

The simulated results are compared with the observed values of outcome variables between 1978 and 1998 to assess the effect of the exogenous variable(s). Unless otherwise indicated, the counter-factual simulations rely on regression estimates for specification 4 reported in Tables 1 and 2. The results are reported in Table 3.

The first counter-factual considers the effects of income on living arrangements and, thus, on income inequality. The simulation reflects the combined effects of greater earnings and changes in the relative earnings of workers and pensioners. The regression estimates are used to "predict" what living arrangements would have been in the absence of changes in the income variables. Comparing the simulation to the actual changes in living arrangements and income inequality during the period leads to the following conclusions. First, changes in income between 1978 and 1998 led to a decline in the proportion of adults living in extended households by 43.5%; a rise in the dependency ratio in extended households by 50.7%; and a decline in the pooling effect by 29.0%. Second, income inequality increased by 2.8% during the period because of the changes in living arrangements induced by changes in income.

Table 3 about here. Counter-factual Analysis

The second counter-factual simulation addresses the effects of changes in age structure on income inequality, incorporating both compositional effects and living arrangement effects. Living arrangements are predicted holding age structure constant at the 1978-level but allowing other independent variables to equal the observed values in each year using specification 2.[11] Income inequality is calculated using these predicted living arrangement variables and the 1978 age distribution. Comparing the counter-factual simulation with the actual changes provides an estimate of the full effect of age structure on income inequality incorporating both compositional effects and living arrangement effects. Population aging led to a 25.3% increase in the proportion living in extended households, a 91.9% increase in the dependency ratio in extended households, and a 59.5 percent increase in the pooling effect. Despite the increased pooling effect, income inequality rose by 2.5% between 1978 and 1998 because of the compositional effects of population aging.

The third and fourth counter-factual simulations presented in Table 3 address the separate effects of fertility decline and improvements in survival. The counter-factual simulation for fertility was carried out by assuming that the "net reproduction rate" did not decline after 1978 but remained at the high level observed in that year – about 2.5 surviving offspring per pensioner for all family cohorts. We adjusted the age-structure to generate the younger adult age structure that would have existed in the absence of fertility

[11] Specification 2 is used in this instance because the dependency ratio captures the average effects of changes in age structure due both to changes in fertility and survival.

decline that occurred after 1978. The effect of fertility decline was to increase the proportion living in extended households by 10.9%, the dependency ratio by 17.9%, and the pooling effect by 14.8%. Despite the favorable changes in living arrangements, income inequality increased by 3.4% because of the effects of fertility decline on age structure.

The counter-factual simulation of survival holds the survival rate, as measured in the analysis of living arrangements, constant at the 1978 level. The age-distribution is adjusted to reflect the lower rates of survival in the counter-factual analysis. Again compared with the actual changes between 1978 and 1998, improvements in survival led to a 10.4% increase in the proportion extended, a 60.5% increase in the dependency ratio and a 28.8% increase in the pooling effect. Reduced survival led to a 1.5% increase in income inequality.

One of the surprising results that emerges from the counter-factual simulations is that improvements in survival have had a stronger pooling effect than have reductions in fertility. The simulated effects reflect both the partial effect of each variable and the magnitude of the change in the variable in question. Survival has a smaller partial effect as determined from the statistical analysis but it changed more during the twenty-one year period analyzed. But improvements in survival led to a much greater increase in the representation of pensioners in extended households. As a result changes in survival had a more favorable pooling effect than changes in fertility.

The results presented in Table 3 provide estimates that incorporate both the compositional effects and the effects on living arrangements of aging. Clearly the response in living arrangements was insufficient to reverse the compositional effect of aging on income inequality, but what would have been the effect of aging had living arrangements not changed? This question is addressed in Figure 4, which compares the full effect of aging on income inequality in each year from 1978 to 1998 with the effect of compositional changes alone. In 1998, income inequality would have been higher by 4.3% rather than 2.5% had aging not induced changes in living arrangements that were more favorable to low income inequality.

Figure 4. Effects of age structure on income inequality

VI. Limitations and further work

The analysis reported here relies on important simplifying assumptions that require qualification of the conclusions. Perhaps the most important is the assumption that co-residence decisions are made independently of income. To the extent that pensioners with atypically low income relative to their children are more likely to form extended households, income inequality will be reduced by more than implied by our simple model. On the other hand, to the extent that lower income *families* are more likely to establish extended households, income inequality will be reduced by less than implied by our simple model. A priority in future work, then, is to analyze a more complete and realistic model of income inequality.

A second problem with the application of the simple model is that it involves an unknown parameter – the correlation between the current incomes of pensioners and workers within extended households. Several studies have estimated the correlation between the earnings of parents and their offspring, at similar stages in their lifecycle, but we are not aware of estimates of the correlation in the current income of workers and their mostly retired parents. Sensitivity analysis shows, however, that the results presented here are insensitive to substantial variation in the correlation coefficient.

There are important aspects of income inequality that are not addressed by this study. We have excluded children and young adults. In Taiwan and many other countries there are important changes in the age at which children enter the labor force, marry, and establish separate households. Previous research has shown that changes in age structure at younger ages have important implications for earnings inequality (Lam 1997; Lam and Levison 1992), but the influence on income inequality incorporating changes in living arrangements has been neglected. Likewise fertility differentials may have important implications for per capita income inequality that are not addressed in this paper. Aging and its implications for old-age support is only one of the important ways in which demographic change bears on income inequality.

Analysis is limited to inequality in current income, whereas inequality in lifetime income may be a preferred welfare measure. Age composition has no implications for inequality in lifetime income, but as von Weizsacker (1995) notes changes in transfers induced by aging do affect lifetime income inequality. Although we do not explore this issue here, the response of family support systems bears on both current and lifetime income inequality.

A troubling aspect of this and other studies of income inequality is the lack of information about the intra-household distribution of resources. Measures of income inequality that are in common use essentially assume that resources are equally shared within the household. To the extent that they are not equally shared a rise in the proportion living in extended households has a smaller effect on income inequality. Thus, this and other studies analyze income inequality as it is measured, not as it would ideally be measured.

The limitations of OLF model used in this study have yet to be firmly established. In undertaking this research we have conducted one form of sensitivity analysis that addresses the robustness of the model. In the results reported in the paper, we use single-year of age and 30 family cohorts in each year. We have also carried out the analysis using three-year age groups and 10 family cohorts and five-year age groups with 6 family cohorts. The empirical results are relatively insensitive to this variation. Some of the issues that arise with the OLF model are discussed in more detail in Mason and Lee (2003), but additional work is needed to assess potential problems that may arise because of the simplifying assumptions that underlie the model.

VII. Conclusions

Analyzing income inequality in Taiwan is a challenging exercise. So many aspects of Taiwan's economy and society have experienced rapid change, but income inequality has remained at low levels, changing modestly in the past few decades. Explaining why something has not changed is a good deal more difficult than explaining why it has.

The analysis in this paper does provide some answers to that difficult question. First, population aging had little effect on income inequality because the unfavorable compositional effects were offset by favorable pooling effects. Second, population aging could have had led to a greater increase in the proportion living in extended households and a stronger pooling effect, but improvements in survival have a weaker effect than fertility decline on the proportion living in extended households. A possible explanation for this is that improvements in health led to an increase in the proportion living independently, partially offsetting the availability effects associated with population aging. Third, higher income in Taiwan could have led to a more shift away from extended households. It did not, in part, because the earnings of the pensioner generations declined relative to the earnings of worker generations partially offsetting the effect of the general increase in earnings on the proportion living in extended households. Moreover, changes in income led to greater age balance in extended households with favorable implications for income inequality. As a result of these offsetting forces, income had a more modest effect on income pooling and income inequality.

The future of income inequality in Taiwan is an interesting issue that we have not systematically explored in this paper. The compositional effects of population aging may become favorable in the coming decades (see Figure 3). As the effects of low fertility take hold, the pooling effect may rise also with favorable implications for inequality. On the other hand, other forces, such as the strengthening of public support systems, may undermine the family support system in Taiwan leading to greater income inequality.

There are aspects of this research that are more general interest. First, we have shown that inequality for family cohorts can be modeled as two additive components – a direct aging effect that captures compositional effects and a pooling effect that captures the manner in which families pool their income by forming multi-generation households. Second, the pooling effect is determined by two features of living arrangements. An increase in the proportion of family members living in extended households and greater age balance in extended households lead to a more favorable pooling effect. Third, on *a priori* grounds the components of the pooling effect should be influenced by population aging. Thus, the effects of population aging are not limited to compositional effects. Fourth, the effect of population aging depends on whether fertility decline or mortality decline is the source of population aging. In many developing countries fertility decline is sufficiently recent that low fertility cohorts are just beginning to reach older ages. As they do the implications of low fertility for family support systems will become increasingly important.

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Table 1. Regression results for proportion living in multi-generation households ($\ln m_x$). Standard errors in parentheses.				
Variable name	1	2	3	4
$\ln D$	**0.231 (0.013)**	**0.374 (0.029)**	**0.254 (0.014)**	**0.417 (0.032)**
$(\ln D)^2$	**-**	**0.032 (0.0060)**	**-**	**0.039 (0.0069)**
$\ln s$	**-**	-	**-0.096 (0.025)**	**-0.253 (0.060)**
$(\ln s)^2$	**-**	-	**-**	**-0.046 (0.013)**
Cohort	**0.025 (0.0013)**	**0.027 (0.0014)**	**0.028 (0.0015)**	**0.028 (0.0015)**
$\ln labwork$	**-0.206 (0.018)**	**-0.260 (0.021)**	**-0.227 (0.019)**	**-0.232 (0.023)**
Incrat	-0.047 (0.042)	-0.038 (0.042)	-0.058 (0.037)	-0.041 (0.041)
Sexratio	**-0.150 (0.029)**	**-0.134 (0.028)**	**-0.103 (0.031)**	**-0.080 (0.031)**
Constant	**-48.92 (2.60)**	**-53.46 (2.67)**	**-54.48 (2.95)**	**-54.80 (2.95)**
N	630	630	630	630
Adjusted R^2	0.954	0.956	0.955	0.957
Notes	Bold face indicates significant at 1% level.			

Table 2. Regression results for dependency ratio in multi-generation households ($\ln D_x$). Standard errors in parentheses.

Variable name	1	2	3	4
$\ln D$	**1.016 (0.022)**	**1.174 (0.049)**	**0.998 (0.024)**	**1.164 (0.054)**
$(\ln D)^2$	**-**	**0.036 (0.010)**	**-**	**0.040 (0.012)**
$\ln s$	**-**	-	0.077* (0.042)	0.211** (0.101)
$(\ln s)^2$	**-**	-	**-**	0.019 (0.022)
Cohort	**-0.044 (0.0022)**	**-0.042 (0.0023)**	**-0.047 (0.0025)**	**-0.045 (0.0025)**
$\ln labwork$	**0.329 (0.030)**	**0.270 (0.034)**	**0.346 (0.032)**	**0.269 (0.038)**
Incrat	**-0.221 (0.070)**	**-0.212 (0.070)**	**-0.213 (0.070)**	**-0.197 (0.070)**
Sexratio	**-0.186 (0.047)**	**-0.168 (0.047)**	**-0.223 (0.051)**	**-0.238 (0.052)**
Constant	**86.98 (4.28)**	**80.99 (4.46)**	**91.43 (4.91)**	**77.48 (7.06)**
N	630	630	630	630
Adjusted R^2	0.980	0.980	0.980	0.981
Notes	Bold face indicates significant at 1% level. Significant at the 10% level (*) or 5% level (**).			

Table 3. Effects of Income and Demographic Variables on Living Arrangements and Income Inequality

	Percentage Change in:			
Factor	Proportion extended	Dependency ratio	Pooling effect	Income inequality
Income	-43.5	50.7	-29.0	1.4
Age Structure	25.3	91.9	59.5	2.5
Fertility	10.9	17.9	14.8	3.4
Survival	10.4	60.5	28.8	1.5

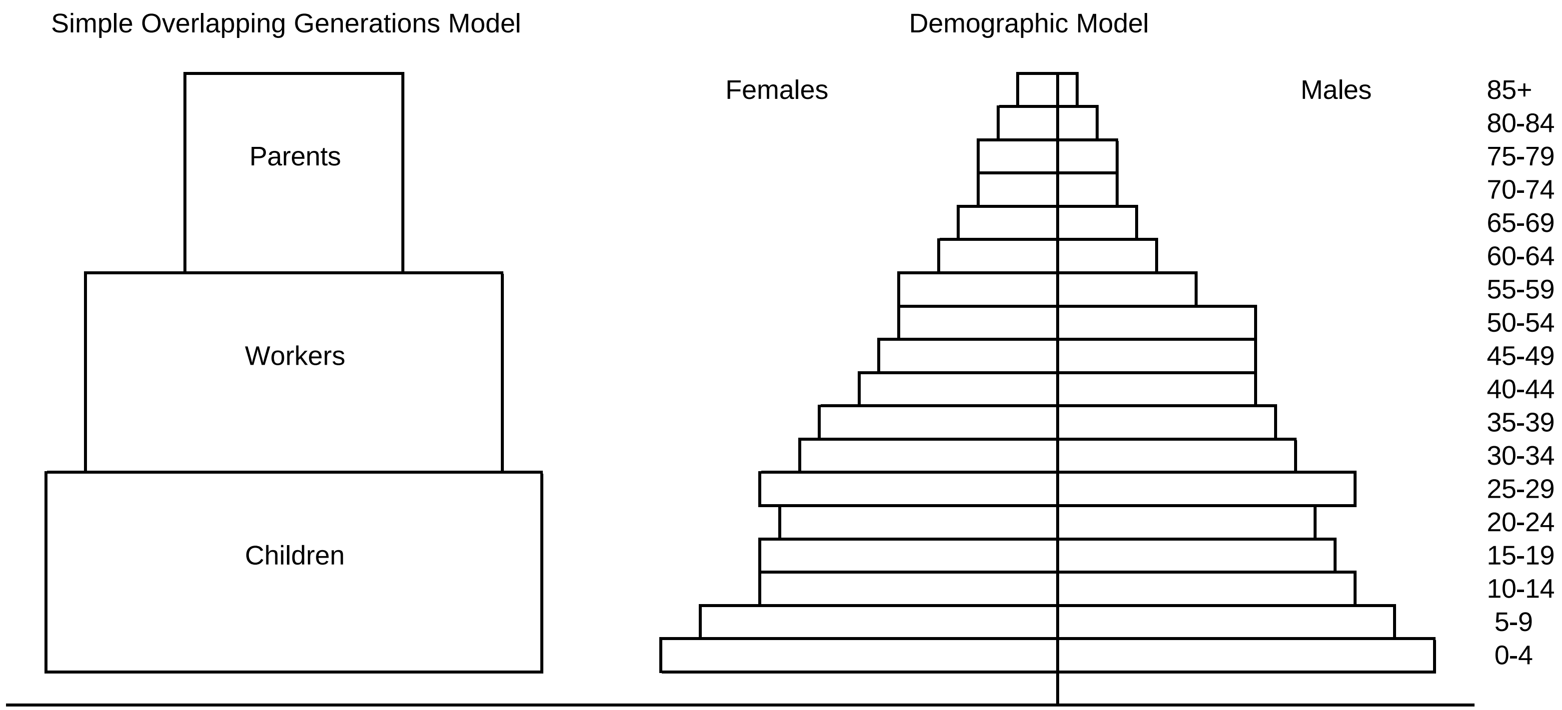

Figure 1. Two Perspectives on Population: Overlapping Generations Model and Demographic Model

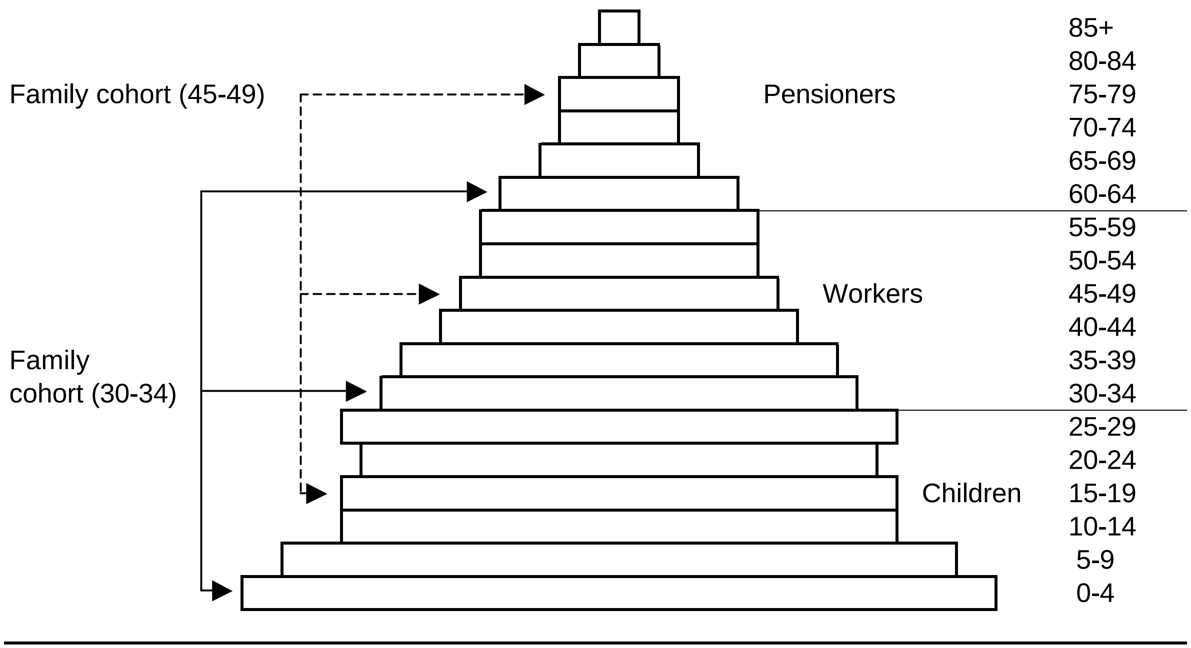

Figure 2. Overlapping Families Model

Figure 3. Aging and Inequality, Family Cohort
V(Y) = 100 for mp = 0.15

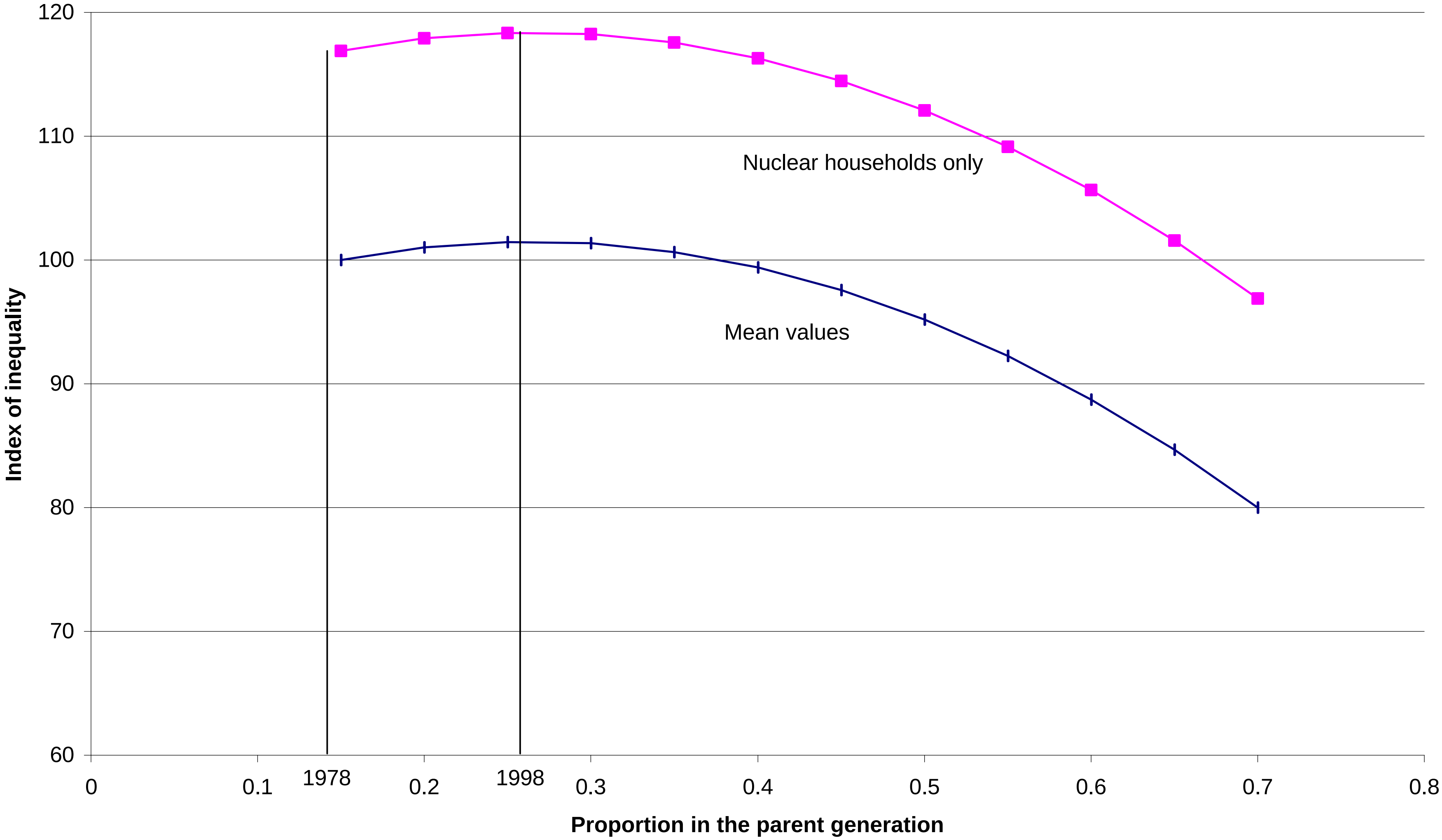

Figure 4. Effects of age structure on income inequality
Percentage difference in CV(Y)