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Population Aging and Living Arrangements: Implications on Inequality†

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Many previous studies 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. These issues are addressed in this paper by (1) developing a new empirical strategy for estimating the effects of population aging, income, and other socio-economic variables on living arrangements; (2) extending the standard model of income inequality to incorporate responses in relation to one aspect of the familial support system – the formation of multi-generational households (or extended households); and (3) applying the models to South Korea, where familial support has been important source of consumption for older people. In particular, this paper complements previous studies by incorporating the responses of familial support systems to changes in the age distribution of the population. Our model and empirical results suggest that (1) population aging could have led to a greater increase in the proportion living in extended households, but improvements in survival have had a weaker effect than the fertility decline on the proportion of people living in extended households, (2) higher incomes of workers in Korea could have led to more of a shift away from extended households, and (3) an increase in both the proportion of the family cohort and the proportion of pensioners living in extended households reduces the variance in income. These results support the argument that co-residence and population aging may reduce income inequality.

Key Word: Population aging, Living arrangement, Extended household
Income inequality, Private transfers
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I. Introduction

Many previous studies have concluded that population aging leads to an increase in consumption or income inequality, which is explained by the permanent income hypothesis (Deaton and Paxson, 1997; Schultz, 1997; Luo *et al*., 2018). These 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). Schultz (1982) devised a decomposition method for analyzing the interactions among income, the number of adults per household, and the number of surviving children. Several other studies have also 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 effects of population aging on income inequality incorporating the responses of public transfer systems have also been explored (von Weizsäcker, 1996).

Depending on the responses of public and private transfer systems, however, analyses limited to compositional effects may be misleading. Some earlier work shows that public transfers can help reduce inequality (Anderson *et al*., 2017; Aaberge *et al*., 2019; Doumbia and Kinda, 2019; Sidek, 2021). Although empirical evidence that private transfers reduce inequality remains insufficient, Lee and Mason (2011) note the role of private transfers in reducing inequality.[1] The prospect of population aging has motivated reforms of many public transfer systems in Latin America, Europe, and the U.S. Although the ultimate effect of reforms on the corresponding redistributive features remains uncertain, Gruber and Wise (2001) conclude that population aging has led to a decline in the share of resources going to the elderly. They estimate that in OECD countries, an increase in the share of the elderly population by 1 percent leads to a rise in spending on the elderly by 0.47 percent. Similarly, Razin, Sadka, and Swagel (2002) conclude that a rise in the overall dependency ratio is leading to a decline in social transfers. In contrast, Preston (1984) argues that the elderly in the U.S. may claim a disproportionate share of public resources as their numbers and political power increase.[2]

In many countries, the familial support system is as important as 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 (parents) increases and the number of workers (offspring) declines within any family, either workers (offspring) must increase the share of their resources devoted to supporting pensioners (parents) (higher "taxes") or pensioners (parents) must experience a decline in the share of their resources derived from relying on workers (offspring) (reduced benefits). On theoretical grounds, however, the effect of population 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.

[1]Hammer and Prskawetz (2022) argue that private transfers are difficult to capture because they mainly occur through informal channels such as intra-family transactions.

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

However, if altruism is weak or intergenerational transfers are motivated by non-altruistic concerns, income differences 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 rates improve, providing familial support to the elderly becomes a responsibility shared by a larger share of workers. 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 workers 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) developing a new empirical strategy for estimating the effects of population aging, income, and other socio-economic variables on living arrangements; (2) extending the standard model of income inequality to incorporate the responses in one aspect of the familial support system – the formation of multi-generational households (or extended households); and (3) applying the models to South Korea (hereafter Korea), where familial support systems have long been important and where public support systems are also currently being rapidly developed. In particular, this paper complements these previous studies by incorporating the responses of familial support systems to changes in the age distribution of the population.

Building on this framework, the paper develops a theoretical model in which income inequality within family cohorts can be decomposed into two additive components – a compositional effect associated with population aging and a pooling effect that reflects the extent to which families share income through the formation of extended households. The model predicts that population aging influences not only the demographic composition of family cohorts but also the mechanisms through which income is pooled within households. Importantly, the magnitude and direction of these effects depend on whether population aging arises primarily from the decline in fertility or from the mortality improvement, as each process alters the structure and functioning of familial support systems in distinct ways.

An empirical analysis using Korean data provides evidence consistent with these theoretical predictions. Population aging, driven mainly by the fertility decline, tends to increase the proportion of individuals living in extended households, thereby reducing income inequality through enhanced intra-family pooling. However, recent trends away from extended households suggest a weakening of the inequality-reducing benefits of shared living arrangements. Overall, the findings support the view that co-residence within extended families serves as a stabilizing mechanism that mitigates the inequality pressures typically associated with demographic change.

Although the major emphasis of this paper is methodological, the analysis of inequality in Korea is of interest in its own right. Korea has experienced rapid demographic and economic change. The total fertility rate (TFR) in Korea declined from 4.53 births per woman in 1970 to replacement level by 1983, and it fluctuated at around 1.60 during the 1980s. However, it decreased sharply again, from 1.65 to 1.18, from the early 1990s to the early 2000s. After the early 2000s, the TFR remained between 1.09 and 1.30 until the mid-2010s; however, it fell below 1 (0.98) in 2018, eventually dropping to 0.84 in 2020 and 0.72 in 2023, the lowest level in the world. On the other hand, life expectancy in Korea increased continuously, from 62.3 (men 58.70, women 65.80) in 1970 to 83.5 (men 80.5, women 86.5) in 2020, which is the second-highest among OECD countries. Since 1960, the Korean economy has been one of the most rapidly growing economies in the world. Yet despite these and other equally dramatic social changes, income inequality has been steady and low. Between 1992 and 2009, the Gini coefficient (market income) increased continuously from about 0.25 to 0.32. Nonetheless, the Gini coefficient declined by about 5 percent between 2010 and 2016, the period analyzed in this paper, except for 2016 (Yun, 2017). Understanding how Korea achieved such rapid economic growth while avoided rising income inequality is a crucial policy questions, and this paper suggests that the answer may lie, in part, in the adaptive responses of the familial support system.

The paper is organized as follows. Section 2 models the effects of age structure, income, and other variables on living arrangements, which is useful for dealing with certain economic indicators, such as inequality. Section 3 extends the standard method for analyzing the compositional effect of age structure on income inequality to incorporate the effects of income pooling, which occurs when family members establish multi-generational households. Section 4 uses Korean data to estimate the effects of demographic and economic variables on living arrangements and income inequality. Section 5 concludes the paper and discusses reservations and problems.

II. Living Arrangements and Co-residence Model

A. *Setup: An Overlapping Families Model*

The basic setup is similar to the overlapping families (OLF) model of Mason and Lee (2004).[3] Their model is a synthesis of an overlapping generation (OLG) and demographic models.

The population is readily subdivided into generations. The youngest generation consists of those who have not yet given birth, i.e., those younger than age a. The next youngest generation consists of those who have given birth but whose children have not. They are between the ages of a and $a+g$. Older generations are defined in a similar fashion. Depending on life expectancy and the generation length, the number of generations usefully distinguished could exceed three, but in the population, we find that distinguishing three generations appears to be sufficient in

[3]Mason and Lee (2004) developed a model in which population aging changes the availability of kin with whom intergenerational exchanges of any form can take place. They show that 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.

general.

The population can also be subdivided into mutually exclusive and collectively 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. Thus, the family cohort can consist of workers aged 30-34, pensioners aged 60-64, and children aged 0-4, or workers aged 45-49, pensioners aged 75-79, and children aged 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 this 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.

Our model of co-residence draws on the extensive literature on living arrangements and intergenerational transfers, but the effects of population aging are emphasized. Co-residence provides an efficient means by which to carry 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 enables families to reap benefits from economies of scale 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 often sacrifice their privacy and potentially their control over personal resources. Thus, family living arrangements may change over time as the importance of intergenerational transfers changes, as the consumption of goods and services produced within the household shifts relative to goods and services that are purchased in the marketplace and consumed by individuals, or as the interests of generations converge or diverge.

However, living arrangements are not entirely a reflection of choice. Observed patterns may reflect the availability of kinship (Ruggles, 1987; Wachter, Hammel, and Laslett, 1978; Hanum, Newcombe, and Scott, 2024). In general, multi-generational families can more easily establish multi-generational households. Moreover, the age structure of multi-generational households may be influenced by the age structure of multi-generational families.

There is extensive empirical and theoretical literature on multi-generational living arrangements[4] and on intergenerational transfers[5]. The relationship between income and intergenerational transfers in competing theoretical models is succinctly summarized by Lillard and Willis (1997), whereas Palloni (2001) provides a comprehensive review of the literature on income and living arrangements.

[4]Bachrach (1980), Chevan and Korson (1975), Kobrin (1976), Macunovich *et al*. (1995), Mason and Lee (2004), Michael, Fuchs, and Scott (1980), Soldo (1981), Wister and Burch (1983), Wolf (1995).

[5]Altonji, Hayashi, and Kotlikoff (2000), Becker and Tomes (1976), Costa (1997; 1998), Cox (1987), Ermisch (2003), Frankenberg, Lillard, and Willis (2002), Lillard and Willis (1997), Martin (1989), McGarry and Schoeni (1997), Palloni (2001), Thornton and Lin (1994).

Empirical studies focusing on the West have found a weak relationship between income and living arrangements (Borsch-Supan *et al*., 1992; Michael, Fuchs, and Scott, 1980; Schwartz, Danziger, and Smolensky, 1984), as well as a weak relationship between income and transfers (Altonji, Hayashi, and Kotlikoff, 2000). On the other hand, income was found to have a 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. Income also appears to have a much stronger impact on transfers in developing Asia than in the West (Frankenberg, Lillard and Willis, 2002; Lillard and Willis, 1997).

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

These issues are of particular interest here due to the obvious connections between population aging and the availability of kinship. We employ an analytical approach to demonstrate how population aging affects the kinship group, specifically 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, Hammel, and Laslett, 1978; Wolf, 1995).

B. *Multi-generational Families*

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

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

For any family cohort with no migration, 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 $M(a,t)=f(a,t)M(g,t-a)$, where $f(a,t)$ is similar to the net reproduction rate, e.g., the fertility rate. When dropping the age and year indexes to simplify notation, the age composition of the family cohort is measured by the dependency ratio, $D=M(a+g,t)/M(a,t)=s/f$. The proportion of persons belonging to a multi-generational family is expressed as follows:

$$m_x^F = \frac{\alpha M(a+g,t) + \beta M(a,t)}{M(a+g,t) + M(a,t)} = \frac{\alpha s + \beta f}{s+f} \tag{1}$$

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

As is apparent from the equation of $D = M(a+g,t)/M(a,t) = s/f$, 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 during the working ages will lead to a decline in the proportion of pensioners who have been pre-deceased by their children; i.e. $\partial\alpha / \partial s_k > o$, where s_k is the proportion surviving from birth to age a. Reductions in fertility may lead to an increase in the proportion of pensioners who are childless ($\partial\alpha / \partial f > 0$). However, the fertility effect of childlessness is typically expected to emerge over a longer-term perspective.

There is no obvious reason why the proportion of workers belonging to multi-generational 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., $\beta = s$. This systematically underestimates the proportion of workers with a surviving pensioner in a two-sex population unless the mortalities of mothers and of 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 ϕ is the correlation between the survival of mothers and fathers, then the following holds:[6]

$$\begin{aligned} &\beta = s + s(1-s)(1-\phi) \\ &\frac{\partial \beta}{\partial s} = 1 + (1-2s)(1-\phi) \end{aligned} \tag{2}$$

Lemma #1

The elasticity of β with respect to s is expressed as shown below.

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

$$\eta_\beta = \frac{\partial \beta / \partial s}{\beta / s} = \frac{1 + (1-2s)(1-\phi)}{1 + (1-s)(1-\phi)}$$

If the survival of husbands and wives is positively correlated or independent ($0 \le \phi \le 1$), then an increase in the survival rate leads to an increase in the proportion of workers with surviving parents ($\partial \beta / \partial s \ge 0$). Likewise, the elasticity is positive but less than one ($0 \le \eta_\beta \le 1$).

The partial effect of population aging on the proportion of the family cohort who are members of multi-generational families can now be assessed using equation (1). First, consider the effect of population aging due to an increase in survival.

Lemma #2

Using α' to represent $\partial \alpha / \partial s (>0)$ and β' to represent $\partial \beta / \partial s (>0)$:

$$\frac{\partial m_x^F}{\partial s} = \frac{(\alpha - \beta) f + \alpha'(s+f)s + \beta'(s+f)f}{(s+f)^2}$$

$$\Leftrightarrow \frac{\partial m_x^F}{\partial s}\frac{\partial s}{\partial D} = \frac{(\alpha - \beta) f + \alpha'(s+f)s + \beta'(s+f)f}{(s+f)^2} f$$

If $\alpha > \beta$, population aging due to an increase in survival leads to an increase in the proportion of the family cohort belonging to multi-generational families $\left(\frac{\partial m_x^F}{\partial s} > 0 \; if \; \alpha > \beta \right)$. Also, note that even if $\alpha < \beta$, the partial effect will be positive if the effects of survival on α and β, i.e., α' and β', are sufficiently great. However, if $\alpha < \beta$ and α' and β' are not large, the partial effect would be negative. Also considered is the effect of population aging due to an increase in fertility.

Lemma #3

Letting α'' represent $\partial \alpha / \partial f$:

$$\frac{\partial m_x^F}{\partial f} = \frac{-(\alpha - \beta)s + \alpha''(s+f)s}{(s+f)^2}$$

$$\Leftrightarrow \frac{\partial m_x^F}{\partial f}\frac{\partial f}{\partial D} = \left[\frac{(\alpha - \beta)s - \alpha''(s+f)s}{(s+f)^2} \right] \left[\frac{f^2}{s} \right]$$

If $\alpha > \beta$ and α'' is small, i.e., if fertility has a negligible effect on the rate of childlessness, population aging due to the fertility decline also leads to an increase in the proportion of the family cohort belonging to multi-generational families; i.e., $\frac{\partial m_x^F}{\partial f} < 0$ if $\alpha > \beta$ and α'' is small. However, if the 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-generational families will decline.

C. *Age Structure of Multi-generational Families*

The second availability measure relevant to this analysis is the age structure of multi-generational families, denoted by D_x^F. The relationship between the age structure of the population (D) and the age structure of multi-generational families (D_x^F) is expressed as follows:

$$D_x^F = \frac{\alpha s}{\beta f} = \frac{\alpha}{\beta} D \tag{3}$$

Lemma #4

The partial effect of a change in the age structure of multi-generational families due to a change in the survival rate is given by

$$\frac{\partial D_x^F}{\partial s} = \frac{\alpha}{\beta} \frac{1+\eta_\alpha - \eta_\beta}{f}$$

$$\Leftrightarrow \frac{\partial D_x^F}{\partial s}\frac{\partial s}{\partial D} = \left[\frac{\alpha}{\beta} \frac{1+\eta_\alpha - \eta_\beta}{f}\right] f = \frac{\alpha}{\beta}(1+\eta_\alpha - \eta_\beta)$$

where $\eta_\alpha = \frac{\partial \alpha}{\partial s}\frac{s}{\alpha}$ is the elasticity of α with respect to s, which is greater than zero, and $\eta_\beta = \frac{\partial \beta}{\partial s}\frac{s}{\beta}$ is the elasticity of β with respect to s, which is between zero and one.

In the polar case, i.e., with η_α equal to zero and η_β equal to 1, a change in the survival rate has no effect on the age structure of multi-generational families; i.e., $\frac{\partial D_x^F}{\partial s} = 0$. In more realistic cases, an increase in the dependency rate due to improving survival leads to a rise in the dependency ratio in multi-generational families; i.e., $\frac{\partial D_x^F}{\partial s} > 0$.

Lemma #5

Letting α'' represent $\partial\alpha / \partial f$ and $\eta'_{\alpha} = \frac{\partial\alpha}{\partial f}\frac{f}{\alpha} = \alpha''\frac{f}{\alpha}$ represent the elasticity of α with respect to f, the partial effect of a change in the age structure of multi-generational families due to a change in fertility is expressed as shown below.

$$\frac{\partial D_x^F}{\partial f} = -\frac{\alpha s}{\beta f^2} + \frac{\alpha'' s}{\beta f} = -\frac{\alpha s}{\beta f^2}(1-\eta'_{\alpha})$$
$$\Leftrightarrow \frac{\partial D_x^F}{\partial f}\frac{\partial f}{\partial D} = \frac{\alpha}{\beta} - \frac{\alpha''}{\beta} f = \frac{\alpha}{\beta}(1-\eta'_{\alpha})$$

The partial effect of a change in the age structure of multi-generational families due to a change in fertility is negative if $a''<0$ i.e., $\frac{\partial D_x^F}{\partial f} < 0$ if $\alpha''<0$. However, if the fertility decline is accompanied by a significant rise in childlessness, then this could be positive.

D. *Additional Effects of Survival, Fertility, and Immigration*

The effects of survival and fertility analyzed in the preceding section 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 the longer life expectancy leads to improvements in health status is a complex issue (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 subset to which workers belong. To the extent that multi-generational households consist of only one sibling and his or her pensioners, a common arrangement in East Asia, the fertility decline will increase the probability that a member of the worker generation is living in an extended household.

The simple model presented in this section presumes a closed population. However, international migration may influence the prevalence and age structure of multi-generational 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 migration involves multi-generational families. Second, if immigration is highly age- and sex-specific, the result may be an imbalance in the sex ratio, interference in the formation of marital unions, and a rise in childlessness for members of the surplus sex. Third, marriage migration, one of the most common types of migration in Korea, may also constitute a distinctive form of migration. In the context of Korea, marriage migration typically involves foreign brides entering the country through marriage to local men. Such inflows may not only mitigate local imbalances in the sex ratio but also contribute to the maintenance of multi-generational households by adding potential caregivers or workers within the family. Consequently, marriage migration may partly offset some demographic and familial effects associated with the assumption of a closed population.

III. Co-residence and Income Inequality Model

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

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

Given the mean $(E(Y_K))$ and variance $(V(Y_K))$ of income for family cohorts, income inequality for the population depends only on compositional effects. Applying the conventional formula to family cohort data (Goldstein and Lee, 2014) gives

$$
\begin{aligned}
&V(Y) = \sum_{(k=a)}^{(a+g)} u_k V(Y_k) + \sum_{(k=a)}^{(a+g)} u_k \left\{E(Y) - E(Y_k)\right\}^2 \\
(4)\qquad &E(Y) = \sum_{(k=a)}^{(a+g)} u_k E(Y_k) \\
&CV = \frac{\sqrt{V(Y)}}{E(Y)}
\end{aligned}
$$

where $V(Y)$ is the variance in per adult household income for the population, $E(Y)$ is the mean income per adult, u_k is the proportion of the adult population belonging to the family cohort of age k, and CV is the coefficient of variation. The values are summed over the age interval for the worker generation (from a to $a+g$) because, as noted above, the ages of the worker generation are used to index family cohorts.

Before equation (4) can be employed, however, we must determine how the family cohort variables $V(Y_k)$ and $E(Y_k)$ are influenced by the age composition and living arrangements of family cohorts. Each family cohort consists of two generations, workers aged k and pensioners aged $k+g$ with mean incomes of $E(Y_k^w)$ and $E(Y_k^p)$, respectively, and corresponding variances of $V(Y_k^w)$ and $V(Y_k^p)$. These values are taken as given throughout the analysis. The proportion of family members who are workers is designated by m_k^w and the proportion who are pensioners by m_k^p where $m_k^w + m_k^p = 1 \ (a \leq k < a+g)$.

The mean income of the family cohort, expressed as

$$E(Y_k) = m_k^w E(Y_k^w) + m_k^p E(Y_k^p) \tag{5}$$

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 the age structure and living arrangements. It is this relationship that is key to understanding how changes in living arrangements influence income inequality within the family cohort and the broader population. The remainder of this section analyzes how changes in the age structure and living arrangements influence $V(Y_k)$.

The analysis is greatly eased through a number of simplifying assumptions. 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 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:

$$\begin{aligned} & V(Y_k) = w_1 V(Y_k^w) + w_2 V(Y_k^p) + w_3 C(Y_k^w, Y_k^p) + w_4 \left\{E(Y_k^w) - E(Y_k^p)\right\}^2, \text{ where:} \\ & w_1 = m^w - m_x^w m_x^p m_x^F \\ & w_2 = m^p - m_x^w m_x^p m_x^F \\ & w_3 = 2 m_x^w m_x^p m_x^F \\ & w_4 = m^w m^p - m_x^w m_x^p m_x^F \end{aligned} \tag{6}$$

The coefficients, w_i, and the variables that determine them all vary across family cohorts, but k 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^w) and pensioners (m^p); the age structure of extended households – the proportion of members who are workers (m_x^w) and pensioners (m_x^p); and the proportion of cohort family members who live in extended households (m_x^F). All four coefficients are non-negative; i.e., $w_i \geq 0$ for $i = 1,2,3,4$. The first three coefficients sum to one. Hence, the variance in per adult income, $V(Y_k)$, is determined, in part, as the weighted average of the variances of the incomes of the family members, $V(Y_k^w)$ and $V(Y_k^p)$, and the covariance between their incomes, $C\left(Y_k^w, Y_k^p\right)$ the difference-in-variances component. The fourth right-hand-side term, $\left\{E(Y_k^w) - E(Y_k^p)\right\}^2$, captures the effect of differences in the average incomes of workers and parents, 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 balanced age distribution leads to an increase in the difference-in-means component.

Lemma #6

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:

$$\frac{\partial V(Y_k)}{\partial m^p} = V(Y_k^p) - V(Y_k^w) + (1 - 2m^p)\left\{E(Y_k^w) - E(Y_k^p)\right\}^2$$

The effect of living arrangements is captured by the multiplicative term $m_x^w m_x^p m_x^F$ in equation (6), referred to here as the pooling effect. The pooling effect increases with the age balance in extended household membership $(m_x^w m_x^p)$, reaching a maximum when there are equal numbers of workers and pensioners, and also increases with the proportion of family cohort members who live in extended households (m_x^F).

Defining $R = V(Y_k^p) / V(Y_k^w)$ and ρ as the correlation between the income of parents and workers, equation (6) can be rewritten as follows:

$$V(Y_k) = \left\{w_1 + w_2 R + w_3 \rho \sqrt{R}\right\} V(Y_k^w) + w_4 \left\{E(Y_k^w) - E(Y_k^p)\right\}^2 \tag{7}$$

Lemma #7

The partial effect of an increase in the proportion living in extended households is expressed as shown below.

$$\begin{aligned}\frac{\partial V(Y_k)}{\partial m_x^F} &= m_x^w m_x^p (2\rho\sqrt{R} - 1 - R) V(Y_k^w) - m_x^w m_x^p \left\{E(Y_k^w) - E(Y_k^p)\right\}^2 \\ &= m_x^w m_x^p \left[(2\rho\sqrt{R} - 1 - R) V(Y_k^w) - \left\{E(Y_k^w) - E(Y_k^p)\right\}^2\right]\end{aligned}$$

If $\rho = R = 1$, then $2\rho\sqrt{R} - 1 - R$ is zero; otherwise, it is always negative.[7] Hence, the partial effect of an increase in the proportion of the family cohort living in extended households reduces the variance in income; i.e., $\frac{\partial V(Y_k)}{\partial m_x^F} \leq 0$.

Lemma #8

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

$$\frac{\partial V(Y_k)}{\partial m_x^p} = (1 - 2m_x^p)m_x^F (2\rho\sqrt{R} - 1 - R)V(Y_k^w) - (1 - 2m_x^p)m_x^F \left\{E(Y_k^w) - E(Y_k^p)\right\}^2$$
$$= (1 - 2m_x^p)m_x^F \left[(2\rho\sqrt{R} - 1 - R)V(Y_k^w) - \left\{E(Y_k^w) - E(Y_k^p)\right\}^2\right]$$

If the share of pensioners is less than one-half, the partial effect is unambiguously negative, as noted earlier, $2\rho\sqrt{R} - 1 - R \leq 0$. Hence, $\partial V(Y_k)/\partial m_x^p < 0$ for $m_x^p < 0.5$. As an empirical matter in Korea, the proportion of pensioners who live in extended households (m_x^p) is very low. Hence, we can expect with some certainty that the partial effect of an increase in the proportion of pensioners in extended households reduces the variance in income; i.e., $\frac{\partial V(Y_k)}{\partial m_x^p} \leq 0$. However, as the proportion of cohort family members who live in extended households (m_x^F) is rapidly decreasing, the effect would be smaller now than the past.

IV. Empirical Analysis

The analysis of this research consists of two distinct elements. First, we estimate the effects of population aging and other variables on living arrangements. Second, we estimate the effects of living arrangements and population aging on income inequality.

A. *Data*

Korea is an interesting case study given the country's rapid population aging and rapidly changing living arrangements. As previously mentioned, the total fertility rate (TFR) in Korea declined from 4.53 births per woman in 1970 to 0.72 in 2023, the lowest level in the world. Simultaneously, life expectancy in Korea has continuously increased, going from 62.3 in 1970 to 83.5 in 2020. As a result, Korea is experiencing a severe population aging crisis. In addition, there is a strong

[7] $2\rho\sqrt{R} - 1 - R = -(\rho - \sqrt{R})^2 - (1 - \rho^2) \Leftrightarrow$ If $\rho = R = 1$, then $2\rho\sqrt{R} - 1 - R$ is zero; otherwise, it is always negative because $(\rho - \sqrt{R})^2 \geq 0$ and $\rho \leq 1$.

responsibility for filial piety, known as "Hyo" in Korean, and there has been an obligation for adult children to live with their parents in many Asian cultures, including Korea (Chen *et al*., 2022; Chu *et al*., 2011; Zhang *et al*., 2014). In practice, the multi-generational household system had been relatively widespread in Korea until recently, and it is dramatically shrinking. According to census data from the Ministry of Data and Statistics, the proportion of two-generational households merely consisting of "a couple and parents" or "grandparents and unmarried grandchildren," three-generational households, and households consisting of four or more generations in all households changed from about 12.47% in 2000 to about 8.53% in 2020 for households and from about 18.23% in 2000 to about 13.40% in 2020 for individual household members.

The empirical analysis is constructed using data from the 'Household Income and Expenditure Survey' (HIES) and from the 'National Transfer Accounts' (NTA) for Korea, which were collected annually from 2010 to 2016. Given that the HIES data in Korea was surveyed on a household basis, we use the Korean NTA data to allocate household income to household members. Household income in the HIES data is assigned to each household member based on the age-specific income ratio calculated through the NTA data, considering the age composition of household members in each household.[8] Income is measured by total labor income, which is calculated by summing earnings (wage income) and business income (self-employed income). The means and variances of income are all based on income per adult household member.

The primary dataset consists of values for single-year family cohorts spanning seven consecutive years. The generation length g is set to 30 years. This estimate is based on the conventional difference in the average ages of different generations observed in general. The data are from 30 family cohorts for each year consisting of workers aged 26-55 and pensioners aged 56-85, considering data availability in Korea. Those aged 85 and older are included with the 85-year-old pensioners. Thus, the dataset consists of seven years times 30 age groups, yielding 210 observations.

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 the relationship to the household head information. For example, the head's generation consists of all those reporting as head, the spouse of the head, or siblings, including their spouses, of the head. A second generation consists of all those reporting as parents or grandparents of the head or the head's spouse. Any household that includes adult members belonging to two different generations is classified as an extended household. A household consisting of only one generation is classified as a nuclear household. Marital status is not a factor in the classification scheme. Also, young generational cohorts under the age of 25, mainly including children and grandchildren of the head, are excluded from the analysis.[9] Each household member is designated as a worker or pensioner based on his or her age. If a household member is 56 years or older, he or she is classified as a pensioner.

[8]We can calculate income characteristics, including the income distribution of household members based on the age composition, separately for each household using HIES data and NTA data.

[9]The share of household income for the young generation cohort under the age of 25 is relatively small, and furthermore, this cohort does not significantly affect decision-making within the household in general.

Otherwise, i.e., if a household member is 55 years or younger, he or she is classified as a worker. (Table A2 presents summary statistics for single-year family cohort data for seven consecutive years between 2010 and 2016 prepared using the HIES data and NTA data in Korea.)

B. *Results*

1. Living Arrangements

Regression estimates for the proportion of family cohort members living in extended households $(\ln m_x^F)$ are reported in Table 1; for the dependency ratio in extended households $(\ln D_x^F)$, these values are given in Table 2. All regressions are estimated using ordinary least squares. The estimated regression equation takes the following general form:

$$\begin{aligned} \ln y_x^F &= \alpha + \beta_1 \ln D + \beta_2 (\ln D)^2 + \beta_3 \ln s + \beta_4 (\ln s)^2 \\ &\quad + \gamma_1 Cohort + \gamma_2 \ln labwork + \gamma_3 Incrat + \gamma_4 Sexratio + \delta_a + \varepsilon \end{aligned} \tag{8}$$

where y_x^F denotes either m_x^F or D_x^F, depending on the specification.

For each dependent variable, results from four specifications are reported. Specifications 1 and 2 capture the effects of the age structure using the dependency ratio and its square, respectively. Specifications 3 and 4 introduce adult survival into the model in order to analyze the separate effects of fertility and survival.

TABLE 1—REGRESSION RESULTS FOR THE PROPORTION LIVING IN MULTI-GENERATIONAL HOUSEHOLDS ($\ln m_x^F$)

	(1)	(2)	(3)	(4)
$\ln D$	0.084*	0.145**	0.448***	0.366***
	(0.043)	(0.057)	(0.074)	(0.072)
$(\ln D)^2$		0.037		-0.168***
		(0.023)		(0.032)
$\ln s$			-0.444***	-3.079***
			(0.076)	(0.516)
$(\ln s)^2$				0.255***
				(0.053)
Cohort	-0.001	0.002	0.016***	0.008*
	(0.003)	(0.003)	(0.004)	(0.004)
$\ln labwork$	-1.182***	-1.186***	-1.098***	-0.833***
	(0.099)	(0.098)	(0.092)	(0.098)
Incrat	0.215***	0.099	-0.264**	-0.021
	(0.082)	(0.109)	(0.112)	(0.113)
Sexratio	-0.506***	-0.510***	-0.297*	-0.318*
	(0.178)	(0.177)	(0.169)	(0.162)
Constant	17.072***	11.670*	-15.101**	5.339
	(5.753)	(6.668)	(7.670)	(8.491)
N (observation)	210	210	210	210
Adj. R^2	0.8621	0.8631	0.8818	0.8975

Note: 1) These estimates include year dummies; 2) ***, **, and * denote the 1%, 5%, and 10% significance level, respectively; 3) Standard errors are given in parenthesis.

TABLE 2—REGRESSION RESULTS FOR DEPENDENCY RATIO IN MULTI-GENERATIONAL HOUSEHOLDS ($\ln D_x^F$)

	(1)	(2)	(3)	(4)
$\ln D$	1.132***	0.993***	0.770***	0.939***
	(0.058)	(0.078)	(0.105)	(0.102)
$(\ln D)^2$		-0.085***		0.146***
		(0.032)		(0.046)
$\ln s$			0.442***	4.590***
			(0.109)	(0.734)
$(\ln s)^2$				-0.431***
				(0.075)
Cohort	-0.001	-0.008*	-0.018***	-0.002
	(0.004)	(0.005)	(0.006)	(0.006)
$\ln labwork$	0.781***	0.792***	0.698***	0.326**
	(0.135)	(0.133)	(0.132)	(0.140)
Incrat	-0.685***	-0.423***	-0.209	-0.532***
	(0.112)	(0.148)	(0.159)	(0.160)
Sexratio	-0.237	-0.226	-0.444*	-0.243
	(0.244)	(0.240)	(0.240)	(0.231)
Constant	-7.455	4.823	24.548**	-12.892
	(7.887)	(9.039)	(10.937)	(12.083)
N (observation)	210	210	210	210
Adj. R^2	0.8365	0.8414	0.8484	0.8691

Note: 1) These estimates include year dummies; 2) ***, **, and * denote the 1%, 5%, 10% significance level, respectively; 3) Standard errors in parenthesis.

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 (*Sexratio*), and single-year age dummy variables (δ_a). The coefficients of the age dummies are not reported in the table.

The coefficient of the natural log of workers' average earnings $(\ln labwork)$ captures the effect of general increases in earnings because the ratio of pensioners' average earnings to workers' average earnings is controlled. In specification 4, the estimated coefficient of $\ln labwork$ is -0.833, which is statistically significant at the 1% significance level. This magnitude implies that a 1% increase in workers' earnings is associated with approximately a 0.83% decrease in the proportion of family members living in extended households. The statistically significant negative coefficient is consistent with the standard view that higher income leads to an increase in the demand for privacy. Individuals with higher incomes are willing to give up the gains from economies of scale or public goods captured by extended households. A rise in the earnings of pensioners relative to workers (*Incrat*) has no discernible effect on the proportion of family members living in extended households. The estimated coefficients are not consistent according to the specification. Specifications 1 and 2 show a positive effect, whereas specifications 3 and 4 show a negative effect. Thus, we find no support for the altruism hypothesis. Note, however, that current earnings may poorly measure the economic status of the pensioner generation. The effects of pensioners' wealth may be more supportive of the altruism hypothesis, as has been the case in other studies (Borsch-Supan *et al.*, 1996).

Adequate wealth measures by generation are not available in the data.[10] The cohort effect (*Cohort*) is positive (except for specification 1), and the sex ratio of the pensioner generation (*Sexratio*) has a statistically significant negative effect, as anticipated. A lower proportion of cohorts with substantial surplus males were living in extended, multi-generational households.

Based on specifications 1 and 2, we can identify that the dependency ratio has a positive, statistically significant effect on the proportion of the family cohort living in multi-generational households. The elasticity rises gradually with the dependency ratio. This provides support for a central feature of this paper – population aging may reduce income inequality by encouraging relatively more co-residence. Specifications 3 and 4 address whether the source of population aging, declining fertility or mortality, matters. The analysis presented above shows that the effects on the 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.

This 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 a decline in fertility. A rise in the dependency ratio due to a decline in fertility produces an increase in the proportion of family cohort members residing in extended households, although the elasticity declines 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, e.g., the fertility 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 the availability effects of fertility. An example is that changes in health status may be captured by survival rates (Zimmer, Martin and Chang, 2002), although the evidence on how health status affects living arrangements is mixed (Borsch-Supan *et al*., 1992; Palloni, 2001). Thus, one possible interpretation of the coefficients is possible. This result implies that the availability effects of increased survival are much smaller than the availability effects of the decline in fertility.

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

(9)
$$\begin{aligned}\frac{\partial \ln m_x^F}{\partial \ln s} &= (\beta_1 + \beta_3) + 2(\beta_2 + \beta_4)\ln s - 2\beta_2 \ln f \\ &= -2.713 + 0.174 \ln s + 0.336 \ln f\end{aligned}$$

and the elasticity of m_x^F with respect to f is equal to

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

$$\frac{\partial \ln m_x^F}{\partial \ln f} = -\beta_1 - 2\beta_2 \ln s + 2\beta_2 \ln f = -0.366 + 0.336 \ln s - 0.336 \ln f \tag{10}$$

where β_i is the *ith* estimated coefficient of specification 4 in Table 1.

The elasticity of m_x^F with respect to s increases as the survival rate and the fertility rate increase. On the other hand, the elasticity of m_x^F with respect to f increases as the survival rate increases and as the fertility rate decreases. In addition, the elasticity of m_x^F with respect to f is still negative for all values of s and f observed here.[11] The bottom line is that population aging resulting from a decline in fertility leads to an increase in the proportion of people living in extended households. Also, population aging resulting from the rise in survival leads to an increase in the proportion of people living in extended households, although the effect of survival is less significant than the effect of fertility.

The analysis of D_x^F reported in Table 2 uses the same specifications used for m_x^F. The cohort effect (*Cohort*) is negative for all specifications. Workers' average earnings $(\ln labwork)$ have a statistically significant positive effect on the dependency ratio in multigenerational households. In specification 4, the estimated coefficient of $\ln labwork$ is 0.326 and statistically significant at the 5% significance level. This implies that a 1% increase in workers' average earnings is associated with approximately a 0.33% increase in the dependency ratio within extended households. The positive coefficient suggests that higher aggregate income among workers may enable families to sustain a larger number of dependents—particularly elderly members—within multigenerational settings. In contrast, a rise in the earnings of pensioners relative to workers (*Incrat*) leads to a significant (except for specification 3) decline in the dependency ratio in multigenerational households. In specification 4, the coefficient of *Incrat* is -0.532 and statistically significant at the 1% significance level. This magnitude indicates that a 1% increase in the relative earnings of pensioners is associated with roughly a 0.53% decrease in the dependency ratio in multi-generational households, implying that as the elderly become more financially independent, the economic necessity of co-residence and within-household dependency weakens. The negative and highly significant effect is consistent with the notion that greater income security among pensioners reduces intergenerational economic dependence. A large surplus in the male pensioner population leads to a decrease in the proportion of pensioners in extended households (*Sexratio*).

[11]For all values of s and f that we observe, $\frac{\partial \ln m_x^F}{\partial \ln f} < 0$. In the data, we identify that the minimum value of $\frac{\partial \ln m_x^F}{\partial \ln f}$ is -1.082 and the maximum value of $\frac{\partial \ln m_x^F}{\partial \ln f}$ is -0.007. On the other hand, for $\frac{\partial \ln m_x^F}{\partial \ln s}$, we observe that the minimum value of $\frac{\partial \ln m_x^F}{\partial \ln s}$ is -0.147 and the maximum value of $\frac{\partial \ln m_x^F}{\partial \ln s}$ is 0.417. For $\frac{\partial \ln m_x^F}{\partial \ln s}$, there were some cases with negative values, but most of them had positive values. (The mean value of $\frac{\partial \ln m_x^F}{\partial \ln s}$ is positive, i.e., 0.182.)

The elasticity of D_x^F with respect to D is close to one in specifications 1 and 2. This is consistent with a straightforward demographic model of the age composition of extended households.

Similar to equations (9) and (10), noting that $\ln D = \ln s - \ln f$, the elasticity of D_x^F with respect to s is equal to

$$\begin{aligned}\frac{\partial \ln D_x^F}{\partial \ln s} &= (\beta_1 + \beta_3) + 2(\beta_2 + \beta_4)\ln s - 2\beta_2 \ln f \\ &= 5.529 - 0.570 \ln s - 0.292 \ln f\end{aligned} \tag{11}$$

and the elasticity of D_x^F with respect to f is equal to

$$\begin{aligned}\frac{\partial \ln D_x^F}{\partial \ln f} &= -\beta_1 - 2\beta_2 \ln s + 2\beta_2 \ln f \\ &= -0.939 - 0.292 \ln s + 0.292 \ln f\end{aligned} \tag{12}$$

where β_i is the *ith* estimated coefficient of specification 4 in Table 2.

The elasticity of D_x^F with respect to s declines as the survival rate and the fertility rate increase, but it is positive for all values of s and f observed here. Also, the elasticity of D_x^F with respect to f decreases as the survival rate increases and as the fertility rate decreases, but it is also still negative for all values of s and f observed.[12] The bottom line is that population aging due to a decline in fertility or survival increase leads to a rise in the dependency ratio in multi-generational households.

2. Effect of Living Arrangements and Population Aging on Income Inequality

Table 3 presents the regression estimates of the variance of income per adult $(V(Y_k))$ on the proportion of family cohort members (m_x^F) and of pensioners (m_x^p) living in extended households. All regressions are also estimated using ordinary least squares. The estimated regression equation takes the following general form:

$$\begin{aligned}V(Y_k) = \alpha + \beta_1 m_x^F + \beta_2 m_x^p + \gamma_1 Cohort + \gamma_2 \ln labwork \\ + \gamma_3 Incrat + \gamma_4 Sexratio + \delta_a + \varepsilon,\end{aligned} \tag{13}$$

[12]For all values of s and f that we observe, $\frac{\partial \ln D_x^F}{\partial \ln s} > 0$ and $\frac{\partial \ln D_x^F}{\partial \ln f} > 0$. In the data, we observe that the minimum and maximum values of $\frac{\partial \ln D_x^F}{\partial \ln s}$ are 0.420 and 1.725, respectively. Also, we can identify that the minimum maximum values of $\frac{\partial \ln D_x^F}{\partial \ln f}$ is -1.251 and -0.318, respectively.

where $V(Y_k)$ denotes the variance of income per adult.

The results contain five specifications for the variance of income per adult. Specification 1 captures the effects of the proportion of family cohort members living in extended households. Specification 2 captures the effect of the proportion of pensioners living in extended households. Specification 3 introduces comprehensive models in order to analyze the separate effects of living arrangements and population aging. In addition, specifications 4 and 5 are the results of a robustness analysis of Specifications 1 and 3 using the proportion of family cohort numbers estimated by specification 4 of Table 1 $\left(m_x^F\right)$ and the natural log of the old-age dependency ratio $(\ln D)$ instead of the proportion of family cohort numbers (m_x^F), respectively.

In all specifications, similar to Tables 1 and 2, 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 (*Sexratio*), and single-year age dummy variables (δ_a). The coefficients of the age dummies are not reported in the table.

Similar to Tables 1 and 2, the coefficient of the natural log of workers' average earnings $(\ln labwork)$ captures the effect of general increases in earnings because the ratio of pensioners' average earnings to workers' average earnings is controlled. In Table 3, the estimated coefficients of the natural log of workers' average earnings $(\ln labwork)$ is 6.607, which is statistically significant at the 1% level of significance.

TABLE 3—REGRESSION RESULTS FOR THE VARIANCE OF INCOME PER ADULT ($V(Y_k)$) ON THE PROPORTION OF FAMILY COHORT MEMBERS LIVING IN MULTI-GENERATIONAL HOUSEHOLDS (m_x^F) AND THE PROPORTION OF PENSIONERS LIVING IN MULTI-GENERATIONAL HOUSEHOLDS (m_x^P)

Variable	(1)	(2)	(3)	(4)	(5)
m_x^F	-2.738*** (0.876)		-2.732*** (0.878)		
m_x^F				-0.234 (1.504)	-1.005 (1.571)
$\ln D$				0.088 (0.132)	0.403* (0.235)
m_x^P		-0.135 (0.401)	-0.105 (0.393)		-1.168 (0.720)
Cohort	-0.120*** (0.007)	-0.116*** (0.008)	-0.119*** (0.008)	-0.120*** (0.009)	-0.123*** (0.009)
$\ln labwork$	6.101*** (0.327)	6.690*** (0.294)	6.125*** (0.340)	6.590*** (0.426)	6.607*** (0.425)
Incrat	4.079*** (0.241)	3.660*** (0.205)	4.089*** (0.244)	3.600*** (0.309)	3.533*** (0.310)
Sexratio	-0.798 (0.512)	-0.441 (0.511)	-0.803 (0.513)	-0.456 (0.539)	-0.594 (0.543)
Constant	151.559*** (14.146)	133.711*** (15.717)	149.475*** (16.196)	144.449*** (20.720)	150.020*** (20.919)
N (observation)	210	210	210	210	210
Adjust R^2	0.8827	0.8770	0.8821	0.8766	0.8776

Note: 1) The unit of variance of income is trillion; 2) These estimates include year dummies; 3) ***, **, and * denote the 1%, 5%, 10% significance level, respectively; 4) Standard errors in parenthesis.

This implies that a 1% increase in workers' average earnings is associated with a roughly 6.6 unit rise in the variance of income per adult. The magnitude of this effect suggests that overall income growth among the working generation tends to widen within-household and cross-generational income dispersion. Higher worker earnings may amplify the heterogeneity in economic well-being across family cohorts, particularly when some households benefit more than others from labor market gains. Similarly, the estimated coefficient of the ratio of the average earnings of pensioners to workers (*Incrat*) is 3.533, also statistically significant at the 1% level. This result indicates that a 1% increase in pensioners' earnings relative to workers' earnings leads to an approximately 3.5 unit increase in the variance of income per adult. The positive and significant effect implies that as the income of the elderly rises relative to that of workers, income dispersion within the population widens, possibly reflecting divergent income trajectories between pensioner and worker generations. This finding is consistent with the broader evidence that increasing income among pensioners contributes to greater cross-cohort income inequality, as observed in other contexts with aging populations (e.g., Borsch-Supan *et al*., 1996). The cohort effect (*Cohort*) is significantly negative, indicating higher income variance in the working generation than in the pensioner generation. Also, the sex ratio of the pensioner generation (*Sexratio*) has a negative effect, but it is not statistically significant.

Based on specifications 1 and 3, as explained by Lemma #7, we can identify that the partial effect of an increase in the proportion of the family cohort living in extended households reduces the variance in income; i.e., $\partial V(Y_k) / \partial m_x^F \leq 0$. In addition, specifications 2 and 3, as explained by Lemma #8, show that the partial effect of an increase in the proportion of pensioners in extended households reduces the variance in income, i.e., $\partial V(Y_k) / \partial m_x^p < 0$, as we anticipated, although this outcome is not statistically significant. These statistically insignificant results can be accepted when considering how they reflect the Korean situation. These results also more clearly support the main argument of this paper – that co-residence and population aging reduce income inequality.

V. Conclusions

Analyzing income inequality, including in the Korean case, is a challenging exercise. Many aspects of the Korean economy and society have experienced rapid change, but income inequality has remained at relatively low levels, experiencing modest fluctuations over the past few decades. Explaining why something has not changed is significantly more difficult than explaining why it has.

There are aspects of this research that are perhaps of more general interest. First, we have shown that inequality for family cohorts can be modeled as two additive components – a direct population aging effect that captures compositional effects and a pooling effect that captures the manner in which families pool their income by forming multi-generational 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 leads 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 a decline in fertility or a decline in mortality is the primary source of population aging. In many developing countries, the fertility decline is sufficiently recent such that low-fertility cohorts are just beginning to reach older ages. As they do, the implications of low fertility for familial support systems will become increasingly important.

The analysis in this paper does provide some answers to that difficult question. First, population aging may have led to a greater increase in the proportion living in extended households; however, improvements in survival have a weaker effect on the proportion living in extended households compared to the fertility decline. A possible explanation for this is that improvements in health led to an increase in the proportion of living independently, partially offsetting the availability effects associated with population aging. Second, the higher income of workers in Korea could have led to more of a shift away from extended households. This result may be linked to the fact that individuals with higher incomes are willing to give up the gains from economies of scale or public goods captured by extended households. On the other hand, a rise in the earnings of pensioners relative to workers has no discernible effect on the proportion of family members living in extended households. Finally, an increase in the proportion of the family cohort living in extended households reduces the variance in income. Furthermore, an increase in the proportion of pensioners in extended households may reduce the variance in income. These results support the argument that co-residence and population aging may reduce income inequality.

Although we briefly addressed income inequality in Korea in this paper in relation to living arrangements and population aging, the future of income inequality in Korea is still an interesting and broad issue that we should continuously explore. The effects of population aging on income inequality may be ambiguous in the coming decades. As the effects of low fertility take hold, the extended household effect may also rise with favorable implications for inequality. On the other hand, other forces, such as strengthening public support systems, may undermine the familial support system in Korea, leading to greater income inequality.

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 incomes relative to their offspring are more likely to form extended households, income inequality will be reduced by more than that implied by our simple model. On the other hand, as this study confirmed using the Korean case, to the extent that lower-income workers are more likely to establish extended households, income inequality will be reduced by less than that 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 of the current income of workers and their mostly retired parents.

In addition, there are important aspects of income inequality that are not addressed by this study. We have excluded children and young adults from our study. In most countries, including Korea, there are important changes in the age at which offspring enter the labor force, marry, and establish separate households. Previous research has shown that changes in the age structure at younger ages have important implications for income inequality (Lam, 1997; Lam and Levison, 1992). However, the influence of incorporating changes in living arrangements on income inequality has largely been neglected. Likewise, fertility differentials may have important implications for per capita income inequality. It should also be noted that many pensioners live together with or near their adult children, not necessarily for economic reasons, but to help with childcare and other forms of family support. Such intergenerational co-residence motivated by caregiving arrangements may partly confound the interpretation of multi-generational living as a purely economic response to income or demographic changes. While addressing this issue is beyond the scope of our study, it would be meaningful to explore the interplay among caregiving roles, living arrangements, and income distribution in future research. More broadly, population aging and its implications for old-age support are only one of the important ways in which demographic change affects income inequality.

The analysis is limited to inequality in current income, whereas inequality in lifetime income may be a more preferred welfare measure. Age structure has no implications for inequality in lifetime income; however, as von Weizsäcker (1995) notes, changes in transfers induced by population aging do affect lifetime income inequality. Although we do not explore this issue here, the response of familial support systems may affect both current and lifetime income inequality.

APPENDIX

TABLE A1—COMPOSITIONS OF GENERATIONS AND LIVING ARRANGEMENTS

[WORKER'S AGE (YOUNGEST WORKER'S AGE) VS. PENSIONER'S AGE (YOUNGEST PENSIONER'S AGE)]

[TABLES 1 & 2]

By Workers' Age	Nuclear Household	Extended Household	Total Household
Worker's Generation	$(1-\beta)f$	βf	f
Pensioner's Generation	$(1-\alpha)s$	αs	s
	$m_x^F = \dfrac{\alpha s+\beta f}{s+f}$	$D_x^F = \dfrac{\alpha s}{\beta f}$	$D = \dfrac{s}{f}$

[TABLE 3]

By Workers' Age	Nuclear Household	Extended Household	Total Household
Worker's Generation	$(1-\beta)f$	$\beta f \to m_x^k = \dfrac{\beta f}{\alpha s+\beta f}$	$f \to m^k = \dfrac{f}{s+f}$
Pensioner's Generation	$(1-\alpha)s$	$\alpha s \to m_x^p = \dfrac{\alpha s}{\alpha s+\beta f}$	$s \to m^p = \dfrac{s}{s+f}$
	$m_x = m_x^F = \dfrac{\alpha s+\beta f}{s+f}$	$D_x^F = \dfrac{\alpha s}{\beta f}$	$D = \dfrac{s}{f} = \dfrac{m^p}{m^k}$

$$Y = m^k Y^k + m^p Y^p$$
$$Y^k = \beta Y^k + (1-\beta) Y^k$$

TABLE A2—DESCRIPTIVE STATISTICS FOR SINGLE-YEAR FAMILY COHORT DATA FOR SEVEN CONSECUTIVE YEARS BETWEEN 2010 AND 2016

		Observation	Mean	Std. Dev.	Minimum	Maximum
(1)	D	210	0.669	0.582	0.119	2.915
(2)	D_x^F	210	0.935	0.356	0.135	2.473
(3)	m_x^F	210	0.227	0.093	0.120	0.462
(4)	m_x^P	210	0.465	0.103	0.119	0.712
(5)	*Cohort*	210	1972.5	8.905	1955	1990
(6)	$\ln labwork$	210	14.446	0.214	13.736	14.833
(7)	*Incrat*	210	0.309	0.381	0.000	1.489
(8)	*Sexratio*	210	0.399	0.069	0.128	0.539

Note: 1) D indicates the old-age dependency ratio of the total population; 2) D_x^F indicates the old-age dependency ratio of multi-generational families; 3) m_x^F represents the proportion who are members of multi-generational families; 4) m_x^P represents the share of pensioners in multi-generational families; 5) *Cohort* indicates the cohort effect; 6) $\ln labwork$ indicates the natural log of workers' average earnings; 7) *Incrat* represents the earnings of pensioners relative to workers; 8) *Sexratio* indicates the sex ratio of the pensioner generation.

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