the criteria suggested by Matsuura (2002), we exclude households that seem to have reported clearly irrational responses. Specifically, we drop the following five types of households from the empirical analysis: 1) those which responded that there is at least one income earner but reports no earning income at all; 2) those which responded that its head is an employed worker but reports no employment income; 3) those which responded that its head is a self-employed worker but reports no self-employment income; 4) those which responded that there is at least one pensioner in it but reports no pension benefits; 5) those whose disposable income is not positive. Excluding those households, the range of the sample size falls to between 6,321 (in 1983) and 7,873 (in 1989), which is 86-94 percent of the original size.

2-2. Inequality measures

We concentrate on four measures to evaluate income inequality to make the estimation results robust. The first is the Gini coefficient, which is one of the most conventional measures and also used in the MHLW's official reports on income redistribution. The Gini coefficient (GC) is given by

$$GC = \frac{1}{2n^2\mu}\sum_{j=1}^{n}\sum_{i=1}^{n}\left|y_i - y_j\right| = 1 + \frac{1}{n} - \frac{1}{n}\sum_{k=1}^{n}\left(\frac{1}{n\mu}\sum_{i=1}^{k} y_i\right),$$

where n is the number of the households in the survey, μ is the mean income, y_i is the i-th household's income ($y_1 \leq y_2 \leq \cdots \leq y_i \cdots \leq y_n$), and the value in the parenthesis on the right hand side shows the cumulative distribution of income up to and including y_i. The closer to zero (unity) the coefficient is, the more equally (unequally) income is assessed to be distributed.

The second is the Atkinson index, which is another commonly used measure. With ε (≥ 0) as a parameter referring to relative inequality aversion, the Atkinson index (AI) is given by

$$AI = 1 - \frac{1}{\mu}\left(\frac{1}{n}\sum_{i=1}^{n} y_i^{1-\varepsilon}\right)^{1/(1-\varepsilon)}, \quad 0 < \varepsilon, \varepsilon \neq 1; \quad AI = 1 - \frac{1}{\mu}\exp\left(\frac{1}{n}\sum_{i=1}^{n}\log y_i\right), \quad \varepsilon = 1.$$

The Atkinson index is linked to a specific form of the social welfare function

$$W = \sum_{i=1}^{n} \frac{y_i^{1-\varepsilon} - 1}{1-\varepsilon}, 0 < \varepsilon, \varepsilon \neq 1; \; W = \sum_{i=1}^{n} \log y_i, \; \varepsilon = 1,$$

and gauges a loss of social welfare due to uneven income distribution. We consider two cases of $\varepsilon = 0.5$ and $\varepsilon = 1$ as is done in the LIS.

The third measure is the mean logarithmic deviation (MLD), which is defined as

$$MLD = \frac{1}{n}\sum_{i=1}^{n} \ln\left(\frac{\bar{y}}{y_i}\right) = \ln \bar{y} - \frac{1}{n}\sum_{i=1}^{n} \ln y_i.$$

MLD is equal to zero if income is completely evenly distributed and has a higher value with more unequal distribution. MLD corresponds to the Atkinson index with ε set to be unity such that $AI = 1 - \exp(-MLD)$.

MLD is useful in that it can easily decompose the factors that affect inequality corresponding to each group's characteristics. For example, dividing the society into m age groups with α_g as a share of g-th age group, and denoting each age group's MLD as MLD_g, simple calculations yield

$$MLD = \sum_{g=1}^{m} \alpha_g MLD_g + \sum_{g=1}^{m} \alpha_g \ln\left(\frac{\bar{y}}{y_g}\right),$$

where y_g is the mean income of the g-th age group. The first term on the right hand side corresponds to the within-age inequality and the second term to between-age.

Finally, we use the logarithmic variance, LV, which is defined as

$$LV = \text{var}(\ln y).$$

We can decompose LV in the same manner as in the case of MLD:

$$LV = \sum_{g=1}^{m} \alpha_g LV_g + \sum_{g=1}^{m} \alpha_g \left(\ln y_g - \overline{\ln y}\right)^2,$$

where LV_g is each age group's LV and the first and second terms on the right hand side correspond to the within-age and between-age inequality, respectively.

It is well established that the Gini coefficient is more sensitive to movements around the mean, while the Atkinson index is more sensitive to changes at the extremes of the distribution. Also, the MLD and LV tend to give more weight to the bottom of the distribution. Despite these statistical differences, similar movements in all four indices would give greater confidence that the indicated change in income inequality is not just a statistical distortion. It should be noted, however, that there is sampling error present in these estimates and, therefore, small changes over time may not be statistically significant.

3. Trends in income inequality

3-1. The trend in income inequality during the 1980s and 1990s

We start with overviewing the trend in income inequality during the past two decades. Table 1 shows how the inequality measures, which are described in the previous section, changed during 1980 and 1998 both on a non-equivalized and equivalized income basis and both for initial and disposable income. For the Gini coefficients we also show the MHLW's official figures for non-equivalized income for comparison. The official figures are somewhat lower than our estimates, but there is no significant difference to our figures in both the levels and trends of inequality. In addition, Figure 1 depicts the trends of Gini coefficients of initial and redistributed income (other measures show almost the same trend as seen in this figure).

We can confirm the following facts from the table and figure. First, income inequality of initial income widened substantially between 1980 and 1998. Its Gini coefficient rose 35 percent during that period to 0.449 in 1998 from 0.332 in 1980. Dividing the whole period up into the 1980s (1980-1989) and the 1990s (1989-1998), we find that the pace of rising inequality was faster in the former period – at 2.2 percent in the 1980s compared to 1.1 percent at an annual rate – as already pointed out by several preceding studies. Figure 1, however, shows a sizable jump from 1980 to 1983, and if this period is excluded, the difference in the pace between the 1980s and 1990s is much smaller (0.9 percent point as opposed to 1.1 percent). This may well be cause for skepticism about the comparability of the 1980 data with those in the subsequent survey years, as will be discussed in more detail below.

Second, a rise in the Gini coefficient of disposable income to 0.337 from 0.286 over the whole period (18 percent) was about half of that of initial income. The discrepancy between the trends of initial and disposable income suggests that redistribution policies succeeded at least partially in preventing income inequality from widening at least on a macro level. In fact, the Gini coefficient of disposable income declined 24.9 percent from initial income in 1998, much more than the 13.9 percent in 1980 due to taxation and social security policies (see the fourth column in Table 1). A reduction in inequality *per se* is welcome, but it should be noted that population aging tends to automatically raise between-age income transfer through public pensions and other social security programs. This makes it difficult to assess the redistributive effects of social policies.

The two facts mentioned above, which can be observed also in the Atkinson indices, *MLD*, and *LV*, raise a curious question; that is, to what extent is Japan an unequal society compared to other industrialized countries? Although Japan is not contained in the LIS dataset, there have been some empirical studies about cross-country differences in income distribution which tentatively included the Japanese data. A series of research papers published by OECD economists – including Oxley *et al.* (1997), Burniaux *et al.* (1998), and Förster and Pearson (2002) – as well as Nishizaki, Yamada, and Ando (1998) and Yamada and Casey (2002) are recent examples. Their analyses in general show that income inequality in Japan, measured by Gini coefficient and other indices, was around the average of the surveyed OECD countries in the mid-1990s. Their comparisons, however, were based on the National Survey of Family Income and Expenditure (NSFIE; conducted by the Statistics Bureau), which tends to show lower income inequality than the SIR. Our analysis based on the SIR is expected to examine the robustness of their arguments.

Table 2 compares Japan's Gini coefficient and the Atkinson index with $\varepsilon = 0.5$, which are based on the micro data from the SIR, to those of the LIS member countries in the latest survey year (around 2000 in most countries) for equivalized disposable income. The left part of the table reveals that Japan is a relatively uneven country in terms of the Gini coefficient and the Atkinson index, both of which place Japan in the seventh highest rank among twenty-eight countries on the list. Within the OECD member countries, income inequality is wider in Japan than in other countries except for the United States and the United Kingdom.

We also find that Japan has been facing a relatively high pace of widening income inequality over the past two decades. The right part of the table confirms this by comparing the absolute changes at an annual rate of the two inequality measures during the period between the early 1980s and the latest survey year. Excluding the three countries of the Czech Republic, Slovenia, and Russia, which are in a transition from a socialist economy to a market economy, Japan is one country whose pace of widening inequality is higher than in many other countries.

On the whole, we can summarize that income inequality has been widening in Japan during the 1980s and 1990s especially for initial income, and also that both the level of inequality and pace of its widening inequality in terms of disposable income are relatively high among industrialized countries.

3-2 Decomposition of widening income inequality

The most plausible reason to explain the uptrend of income inequality in Japan appears to be population aging. Income inequality tends to widen as age increases, so a higher share by the elderly is expected to exacerbate inequality even if other factors remain unchanged. This section examines the impact of aging on income inequality to obtain policy implications[1].

To address this issue, let us first look at Figure 2, which illustrates how the age pattern of the Gini coefficients of equivalized income changed during 1980-1998. We divide the households into twenty age groups of three-year spans (-19, 20-22, 23-35,..., 71-74, 75+) based on the age of the household head, and calculate the Gini coefficient for each age group. In the case of initial income (shown in the left part of the figure), there is a clear uptrend of the Gini coefficient and its age pattern looks relatively stable for each age group, except for the elderly in 1980. It implies two things. First, it seems reasonable to hypothesize that population aging has been a key driving force for changes in the inequality measures. We will investigate this hypothesis later. Second, the data in 1980 seems to have some statistical problems, and if that is the case, it would be misleading to take 1980 as a base year for long-term comparisons. Therefore, we are inclined to take 1983 rather than 1980 as a base year when examining the long-term trend in income inequality. At the same time, the age pattern of inequality of disposable income (shown in the right part of the figure) also looks stable, while the slope is more moderate than that of initial income. Furthermore, no significant discrepancy between 1980 and the subsequent years is observed.

Figure 3 compares shares of each age group's population across survey years. As clearly seen in the figure, the households whose heads are older raised their shares causing a rightward shift of the curves. In fact, the peak of the 1998 curve was the group aged around 50, which was

[1] We found no significant correlation between the level of income equality measures and the population share of the elderly or between their changes, based on the cross-section regressions for the countries listed in Table 2.

about 10-15 years older than the peak group of the 1983 curve, probably corresponding to the time gap between the two survey years. The share of the people aged 60 and above also showed a substantial increase. Together with the seemingly stable age pattern of the Gini coefficient, the change in the age structure implies that widening inequality of initial income during the past two decades was attributable mostly to population aging. If that is the case, policy implications of the uptrend in income inequality would be more limited than otherwise.

Then, we try to more quantitatively gauge the degree to which population aging can explain a rise in income inequality. Our analysis is based on a modification of the method which was originally suggested by Mookherjee and Shorrocks (1982) and has been applied in a series of OECD studies and other studies of income distribution. To decompose changes of the MLD over time (over periods 0 and 1), we have

$$\Delta MLD = MLD^1 - MLD^0 = \sum_{g=1}^{m} \overline{\alpha}_g \Delta MLD_g + \sum_{g=1}^{m} \overline{\alpha}_g \left(\ln \frac{\overline{y}^{01}}{y_g^1} - \ln \frac{\overline{y}^0}{y_g^0} \right)$$
$$+ \sum_{g=1}^{m} \left[\overline{MLD}_g + \overline{\ln\left(\frac{\overline{y}}{y_g}\right)} \right] \Delta \alpha_g + \sum_{g=1}^{m} \overline{\alpha}_g \left(\ln \frac{\overline{y}^1}{\overline{y}^{01}} - \ln \frac{\overline{y}^0}{\overline{y}^0} \right),$$

where $\overline{y}^{01} = \sum_{g=1}^{m} \alpha_g^0 y_g^1$ is the mean income holding the age structure constant and the bars on $\alpha_g, MLD_g, \text{and} \ln(\overline{y}/y_g)$ refer to their means during the period, respectively. The first term on the right hand side indicates the impact of changes in inequality within each age group keeping the structure of the population constant; the second term indicates the impact of changes in inequality between age groups, with the structure of the population constant; and the sum of the third and fourth terms reflects the demographic effect due to changes in the population structure, keeping both the within-group and between-group components constant. We can also decompose change in LV over time in the same way as for MLD so that

$$\Delta LV = LV^1 - LV^0 = \sum_{g=1}^{m} \overline{\alpha}_g \Delta LV_g + \sum_{g=1}^{m} \overline{\alpha}_g \left[\left(\ln y_g^1 - \overline{\ln y}^{01} \right)^2 - \left(\ln y_g^0 - \overline{\ln y}^{0} \right)^2 \right]$$
$$+ \sum_{g=1}^{m} \left[\overline{LV}_g + \overline{\left(\ln y_g - \overline{\ln y} \right)^2} \right] \Delta \alpha_g + \sum_{g=1}^{m} \overline{\alpha}_g \left[\left(\overline{\ln y}^1 - \overline{\ln y}^{01} \right)^2 - \left(\overline{\ln y}^0 - \overline{\ln y}^0 \right)^2 \right],$$

where $\overline{\ln y}^{01}$ is the mean logarithm of income holding the age structure constant. Each term on the right hand side means the same as that in the case of MLD decomposition.

Table 3 summarizes the decomposition of the trends in the MLD and LV over the period from 1983 to 1998. For the increase of the MLD for initial income (0.287) over the whole period, 56.1 percent was caused by the demographic effect. It confirms that population aging was a key factor of widening inequality during the 1980s and 1990s. However, within-age and between-age effects were not negligible, accounting for 24.6 percent and 19.3 percent of the change, respectively, although their magnitudes were well below that of the demographic effect. We observe almost the same decomposition for the MLD for disposable income. For the LV, the demographic effects also explain about half of the changes. While the relative sizes of within-age and between-age effects are somewhat different from the case of the MLD, it should be noted that the within-age effect was substantial for a rise in the inequality of disposable income. Table 3 also compares the pattern of the decomposition between 1983-89 and 1989-98. The relative impact of the demographic effect somewhat decreased in the 1990s for initial income, while it increased for disposable income. The within-age effect decreases in the 1990s for both initial and disposable income, while the between-age effect showed mixed trends.

3-4 Cohort effects vs. age effects

The analysis of income inequality in the previous sections, based on the micro data from the SIR, has serious limitations, in that they are cross-sectional and not panel data. Throughout one's lifetime, every individual goes through young and old stages, the degree of income inequality and the effects of redistribution policies should be evaluated on a lifetime income basis rather than an annual income basis, as suggested by micro simulations conducted by Nelissen (1998), Coronado, Fullerton, and Glass (2000) and others. The SIR, which is a cross-sectional survey, does not contain any longitudinal information so that we cannot directly address inequality and redistribution issues on a lifetime income basis. Instead, we construct synthetic panel data to investigate whether there is any sign of widening income inequality with younger cohorts. Table 3 confirms that population aging is a key factor of widening income inequality during the past two decades, but it also shows a significant contribution from the within-age effect, which cannot rule out the case that the younger cohorts face wider income inequality.

Using the micro data from the SIR of every third year from 1980 to 1998, we construct the synthetic panel data as follows. First, we calculate the values of the inequality measures of equivalized income, both initial and disposable, for each age group (aged 21-23, 24-25,..., 72-74) in each survey year. Next, we set up a flow of their values for the seven survey years for each cohort. For instance, we connect the inequality measures for age 21-23 in 1980, age 24-26 in 1983,..., age 39-41 in 1998 to roughly trace the age pattern of inequality for the cohort which was 21-23 years old in 1980 (that is, those born in 1957-59). We repeat the same procedure up to the cohort which was 54-56 years old in 1980 (born in 1924-26), to construct the age pattern of each inequality measure for twelve cohorts[1]. With these synthetic panel data, we get information about income inequality for eighteen years (six survey years) for each of twelve cohorts, who appeared in all survey years.

Then, we check whether there is any sign of widening inequality in younger cohorts. For this purpose, we first plot the age curves of the Gini coefficients for both initial and disposable income for twelve cohorts in Figure 4. In this figure the age curve of the younger cohort starts

[1] We do not use the date for people aged 60 and above, to avoid any statistical problem that may be included in the income data of the elderly (see Figure 2).

and ends at a younger age. While there is a clear uptrend of income inequality along with age, as already observed in Figure 2, we cannot clearly identify any cohort effect which should be reflected in an upward (or downward) shift of the age curve. In fact, among previous studies there has been no consensus regarding the existence of the cohort effect for income inequality: Iwamoto (2000) recognized the cohort effect in income during the period from 1989 to 1995 based on the Basic Survey of People's Life, whereas Ohtake and Saito (1998) did not find it during the period from 1979 to 1989 based on the NSFIE. Both of them focused on initial income, and no researcher has analyzed the cohort effect in terms of disposable income.

To capture the cohort effect, we first assume linearity for both cohort and age effects and estimate the following equation:

$$inequality\,(cohort, age) = const. + \alpha * cohort + \beta * age + \varepsilon, \tag{1}$$

where the variable *inequality* (*cohort*, *age*) denotes the inequality measures for a certain combination of cohort and age; the variable *cohort* takes the value 1 for the cohort who was born in 1957-59 (21-23 years old in 1980), and 2 for the next cohort, and so on; the variable *age* takes the value 1 for ages 21-23, and two for ages 23-26, and so on; and ε is an error term. The sample size for the estimation is equal to 84 (=12 cohorts multiplied by seven survey years). As already suggested by Figures 2 and 4, however, there appears to be an upward turn of the age effect at around 60 years old, which is a normal retirement age in Japan, especially in the case of initial income. So we also estimate another version of the equation as

$$inequality\,(cohort, age) = const. + \alpha * cohort + \beta * age + \gamma * dum60 + \delta * dum60 * age + error \tag{2}$$

where the variable *dum60* takes the value one for the people aged 60 and above.

Table 4 summarizes the estimation results for equations (1) and (2) in terms of each inequality measure and for initial and disposable income. Three facts are noteworthy with regard to this table. First, equation (2), which takes into account an upward turn at age 60 shows a better fit than equation (1). Second, we cannot find any significant cohort effect for initial income, while we observe a substantial age effect in all cases. It is in line with the results reported by Ohtake and Saito (1998), who focused on the LV of initial income in the NSFIE. Third, when it comes to disposable income, significant cohort effects for seven equations out of ten indicate that the younger cohort faces wider income inequality.

Another way to grasp the cohort effect as well as the age effect is to estimate the equation

$$inequality\,(cohort, age) = const. + \sum_{i=2}^{11} \alpha_i * cohort_i + \sum_{j=2}^{18} \beta_j * age_j + error, \tag{3}$$

where the valuable $cohort_i$ takes the value 1 for the i-th oldest cohort (born in 1924+3i to 1927+3i) and 0 otherwise, and the valuable age_i takes the value 1 for the j-th oldest age group (18+3j to 21+3j years old) and 0 otherwise. We assess the cohort and age effects by estimating the coefficients α_i and β_i, respectively. Table 5 shows the estimation results for disposable income. Generally in line with the results reported in the right part (Equation (2)) of Table 4, most of the coefficients on age dummies are quite significant and those values increase with increasing age, indicating the existence of substantial age effects. More importantly, most of the coefficients on cohort dummies also are significant in the cases of the Atkinson index with $\varepsilon = 0.5$, MLD, and LV, and their values tend to increase as cohort age decreases.

In summary, we cannot deny that the cohort effect has been significant for disposable income during the past two decades in terms of at least some inequality measures, even if population aging has been a main driving force of widening income inequality. In addition, the fact that the cohort effect has been significant for disposable income rather than initial income implies that the current schemes of redistribution policies have failed to fully offset the widening trend of income inequality within the same cohort.

The next issue to be discussed should be the effect of redistribution policies such as taxes and social security schemes. As indicated by Table 1 and Figure 1, income inequality has not widened for disposable income as much as for initial income, implying that redistribution policies have succeeded in enhancing equity at least on an annual income basis. However, the existence of the cohort effect in disposable income, which has been observed in this section, suggests that we should carefully assess the effect of redistribution policies.

4. Redistribution policies

4.1 Decomposition of redistribution

In this section, we examine the effects of redistribution policies. Ideally, we should directly assess the extent to which they have succeeded in reducing inequality of lifetime income within the same cohort. Given the lack of longitudinal information, however, all we can do is to examine the policy impact on within-age inequality and obtain any implications to within-cohort inequality and its trend. For this purpose, we decompose the effects of redistribution policies to between-age and within-age effects in terms of MLD and LV. The more the within-age effects turn out to be limited, the more we have to be cautious in assessing the impact of distribution policies on lifetime income inequality.

A conventional way of decomposing the effect of redistribution policies in terms of MLD in a certain survey year is to take

$$\Delta MLD = \sum_{g=1}^{m} \alpha_g \left(MLD_g^{dis} - MLD_g^{in} \right) + \sum_{g=1}^{m} \alpha_g \left[\ln\left(\frac{\bar{y}^{dis}}{y_g^{dis}} \right) - \ln\left(\frac{\bar{y}^{in}}{y_g^{in}} \right) \right],$$

where suffixes *in* and *dis* denote initial and disposable income, respectively[1]. The first term of the right hand side refers to the within-age effect and the second term the between-age effect. This decomposition, however, could be misleading since the within-age effect is affected by income transfer across age groups. For the elderly, net income transfer from the young – mainly

[1] Ohtake and Saito (1999) applied this method of decomposition in terms of LV. Conceptually, the mean of disposable income ($\bar{y}^{in}$) and the mean of initial income ($\bar{y}^{dis}$) should be the same. However, they are not actually the same due to statistical errors and institutional factors.

through public pensions as well as medical and nursing care programs – tends to raise their mean income, causing a reduction in the within-age MLD, even without any within-age income redistribution. For the young, on the contrary, net income transfer reduces their mean income and increases their within-age MLD. In this sense, the decomposition mentioned tends to overestimate the within-age effect for the elderly and underestimate it for the young.

To avoid this bias, we divide the overall within-age effect into two components: the component which is caused by between-age income transfer and the "pure" within-age effect which is caused solely by income redistribution within the age group (within the sum of initial income and net income transfer receipts). The first component is calculated for the g-th age group as

$$\frac{1}{n_g}\sum_{i\in g}\ln\left(\frac{y_g^{dis}}{y_i^{in}+y_g^{dis}-y_g^{in}}\right)-\frac{1}{n_g}\sum_{i\in g}\ln\left(\frac{y_g^{in}}{y_i^{in}}\right),$$

which grasps the change in the within-age inequality assuming that each household receives the difference between the mean of disposable income (y_g^{dis}) and the mean of initial income (y_g^{in}) for this age group[1]. This component would probably be negative (positive) and indicate that income transfer reduces (raises) income inequality for the elderly (young). The "pure" within-age effect is obtained by subtracting this component from the conventional, overall within-age effect, assuming that the government redistributes the sum of initial income and net income transfer receipts among the age group. Also, we define the sum of the within-age effect caused by between-age income transfer and the conventionally-defined between-age effect as the "total" between-age effect.

We can apply the same decomposition to LV to the within-age and between-age effects in such a way that

[1] Note that the mean of the denominator of the first term across households which belong to the g-th age group is equal to the mean of disposable income for that age group.

$$\Delta LV = \sum_{g=1}^{m} \alpha_g \left(LV_g^{disp} - LV_g^{in} \right) + \sum_{g=1}^{m} \alpha_g \left[\left(\ln y_g^{dis} - \overline{\ln y}^{dis} \right)^2 - \left(\ln y_g^{in} - \overline{\ln y}^{in} \right)^2 \right]$$

and that the component of the within-age effect caused by between-age income transfer for the g-th age group is given by

$$\frac{1}{n_g} \sum_{i \in g} \left[\ln\left(y_i + y_g^{dis} - y_g^{in} \right) - \overline{\ln y_g}^{dis} \right]^2 - \frac{1}{n_g} \sum_{i \in g} \left(\ln y_i^{in} - \overline{\ln y_g}^{in} \right)^2 .$$

4.2 Results

Table 6 summarizes the results of decomposing effects of redistribution policies in terms of the MLD and LV in 1983, 1989, and 1998 for the society as a whole. At first glance, the within-age effect dominated overall income redistribution: it accounts for 85-93 percent of reduction in the MLD and for 56-72 percent of reduction in the LV from initial income to disposable income. However, decomposing the within-age effect into the component reflecting income transfer across the age groups and the pure component reveals that a substantial portion of the within-age effect was due to the former component. Adding it to the conventionally-defined between-age effect, the total between-age effect accounted for about 90 percent of the overall income redistribution in 1998 in terms of both the MLD and LV (see the rightmost column in the table).

Moreover, both the absolute and relative magnitudes of the pure within-age effect kept declining from 1983 to 1998. In the case of the MLD, the contribution of the pure within-age effect was 0.06 points (12.6 percent of the total reduction) in 1998, smaller than 0.08 points (36.7 percent) in 1989, and it was the same case for the LV. In contrast, the total between-age effect increased in importance over the same period. A key factor to explain these changes should be an increasing role played by between-age income transfer due to social security programs under population aging.

It is also interesting to see the age pattern of the decomposition of income redistribution. Figure 5 divides the whole population into three age groups (aged 41 and below, 42-59, and 60 and above) and shows how each redistribution effect contributed to a reduction in the overall MLD by age group in 1983 and 1998[1]. Three facts should be noted from this figure. First, reduction in income inequality concentrated on the elderly in both years: on a net basis, the sum of a reduction in within-age inequality among the elderly and a reduction in the gap between their mean income and the overall mean income accounted more than the entire reduction in income inequality for the society as a whole. As discussed in the previous subsection, however, a reduction in within-age inequality among the elderly was largely caused by income transfer from the young and Figure 5 underlines this.

Second, the between-age income transfer has become more important for overall redistribution of annual income during the past two decades. Indeed, the bar (middle) indicating the within-age effect due to between-age income transfer as well as the bar (right) indicating the between-age effect moved upward for the two young groups and downward for the elderly. To sustain this type of income redistribution, the working generation will likely be forced to pay more taxes. It is also expected to add to within-age inequality due to lower mean income. It is uncertain whether such an increase in between-age income redistribution can improve equity on an annual income basis within the same cohort.

Third, compared to income redistribution caused by between-age income transfer, pure within-age income redistribution was quite limited for the elderly. Indeed, it worked *regressively* for them in 1998, even if its magnitude was very small. Under the current pension programs, the earnings-related component does not contribute to a reduction in income inequality if excluding its effect on a rise in mean income. Also, the current scheme of income taxation, which allows several income deductions for the elderly, tends to work less progressively toward the elderly

[1] The same age pattern is observed for LV.

than the young. The limited magnitude of the pure within-age effect should be taken seriously in discussing redistribution policies, given the uptrend of between-age income inequality and the existence of the cohort effect.

5. Conclusion

We have overviewed the long-term trend of income inequality and the effects of redistribution policies during the 1980s and 1990s in Japan, based on the micro data from the SIR. The key facts that we have confirmed from our empirical analysis are summarized as follows. First, Japan is a relatively uneven society, at least on an annual income basis, and the pace of widening inequality has been faster than in many other countries. Second, while widening inequality during the past two decades was largely attributable to population aging, within-age inequality has also been widening, especially among the elderly, and younger cohorts tend to face wider income inequality of disposable income. Third, income redistribution has been concentrated heavily on between-age redistribution, which accounted for nearly nine-tenths of a reduction in income inequality from initial income to disposable income. Within-age income redistribution has been quite limited in general, if excluding the impact of income transfer from the young.

These facts may well depend much on which survey we choose, and we have to keep in mind the upward bias of income inequality observed in the SIR. However, the estimation results in this paper suggest that we should take seriously the uptrend of inequality measures, even if a substantial portion of it can be explained by population aging. An increasing magnitude of between-age income redistribution via public pensions and other social security schemes also tends to disguise limited effects of the within-age redistribution, allowing for the uptrend of within-age inequality and cohort effects of inequality in disposable income. Limited progressivity of within-age income redistribution – or even regressivity among the elderly – would fail to prevent income inequality within the same age group as well as within the same cohort from rising further.

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Table 1 Income inequality measures: 1980-1998

	Equivalized income			Non-equivalized income	
	Initial (a)	Disposable (b)	(b)/(a)-1	Initial	Disposable
Gini					
1980	0.333	0.286	-13.9%	0.348	0.312
1983	0.383	0.320	-16.4%	0.391	0.342
1986	0.388	0.309	-20.2%	0.402	0.339
1989	0.405	0.327	-19.4%	0.422	0.360
1992	0.413	0.323	-21.7%	0.433	0.362
1995	0.418	0.321	-23.2%	0.434	0.356
1998	0.449	0.337	-24.9%	0.465	0.374
				cf. Gini released by MHLW	
1980				0.349	0.332
1983				0.398	0.358
1986				0.405	0.356
1989				0.433	0.364
1992				0.439	0.365
1995				0.441	0.361
1998				0.472	0.381
Atkinson (ε=0.5)					
1980	0.100	0.068	-32.3%	0.109	0.081
1983	0.145	0.088	-39.6%	0.150	0.099
1986	0.149	0.080	-46.8%	0.159	0.095
1989	0.166	0.091	-45.0%	0.179	0.111
1992	0.174	0.090	-48.6%	0.189	0.112
1995	0.181	0.088	-51.3%	0.191	0.107
1998	0.208	0.097	-53.0%	0.221	0.120
Atkinson (ε=1)					
1980	0.209	0.130	-37.7%	0.232	0.159
1983	0.331	0.165	-50.3%	0.346	0.193
1986	0.358	0.152	-57.7%	0.377	0.186
1989	0.396	0.177	-55.4%	0.419	0.219
1992	0.422	0.174	-58.7%	0.447	0.222
1995	0.441	0.171	-61.3%	0.460	0.211
1998	0.496	0.188	-62.2%	0.516	0.233
MLD					
1980	0.234	0.139	-40.5%	0.263	0.173
1983	0.402	0.180	-55.3%	0.424	0.214
1986	0.443	0.164	-62.9%	0.474	0.206
1989	0.505	0.195	-61.5%	0.543	0.247
1992	0.549	0.192	-65.1%	0.592	0.250
1995	0.581	0.187	-67.8%	0.616	0.237
1998	0.686	0.208	-69.7%	0.725	0.265
LV					
1980	0.634	0.280	-55.8%	0.730	0.372
1983	1.393	0.356	-74.4%	1.472	0.457
1986	1.638	0.329	-79.9%	1.724	0.439
1989	1.891	0.418	-77.9%	1.992	0.563
1992	2.100	0.415	-80.3%	2.214	0.577
1995	2.260	0.399	-82.3%	2.353	0.533
1998	2.661	0.441	-83.4%	2.748	0.588

Source: The author's calculations based on the Surveys on Income Redistribution.