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INCOME INEQUALITY IN TAIWAN 1976-1995: CHANGING FAMILY COMPOSITION, AGING, AND FEMALE LABOR FORCE PARTICIPATION

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Abstract

Change in income inequality in Taiwan from 1964 to 1995 is sensitive to how household incomes are adjusted for household composition. The reasonable practice of dividing household income by persons (or adults) in the household eliminates the widely noted increase in income inequality from 1980 to 1995, and calls into question whether income inequality decreased substantially from 1964 to 1975. The increasing share of the population over age 30 that is associated with the demographic transition has contributed only slightly to increasing income inequality across all ages. The entry of women into the labor force is concentrated among higher wage groups, and thus when one attributes a shadow wage to the time of all persons, regardless of how much they work in the labor force, this broader measure of "full income" inequality is more equal than market income inequality, and it has decreased over time.

JEL Classification: O15, J12
Keywords: Income Inequality, Taiwan, Full Income, Family Composition

1. Introduction

Taiwan is noted as a country that has grown rapidly without increasing income inequality (Fei, Ranis and Kuo, 1979). Taiwan is also one of the first countries outside of the OECD to experience a sharp and sustained decline in its birth rates and is therefore already experiencing an increase in the share of its population older than 30, due to its demographic transition.[1] Primary education was widespread before Japanese rule was terminated in 1945, and expansion of secondary and technical tertiary education proceeded rapidly thereafter, with the difference between the education of men and women narrowing. The labor force participation of women outside of the home has risen while the share of all employment in agriculture fell (Levenson, 1997). This paper considers how the personal distribution of income has changed in Taiwan from 1964 to 1995, viewed against this background of dramatic changes in economic and demographic conditions. It then assesses how changes in the age composition of families, the educational endowments of its people, and the time they supply to the labor force, have contributed to the observed income inequality.

The paper is organized as followed. The next section illustrates the basic dilemma of comparing income inequality across households when the composition of households differs by income level and the composition itself responds to the changing economic opportunities available to adults over their lifetime. Section 3 discusses additional issues that arise in the measurement of income inequality. Section 4 describes the survey data examined in this paper and reviews trends and developments in Taiwan that may be relevant. Section 5 reports changes in household income inequality from 1976 to 1995 in Taiwan and how adult incomes are related to different dimensions of household composition: adult size, proportion of children, and proportion of elderly members. Section 6 decomposes inequality by age of household head to assess the contribution to trends in inequality that are associated with secular changes in age composition of heads of households in Taiwan in this period. Section 7 constructs measures of "full income" inequality that are independent of the changing time allocation of women during this period. The concluding section summarizes the findings.

2. Income Inequality and per Capita Income Inequality: Historical Overview

It is difficult to draw welfare conclusions from changes in household income data if the composition of households is changing at different income levels. Previous studies of the household income distribution in Taiwan illustrate this point; Table 1 tabulates income by household size for 1966, 1972, 1976 and 1995. If the share of income in column (3) is divided by the share of households in column (1), this measure of average household income increases with household size (plotted in Figure 1). But the rate of increase in household income is less than proportional to size (i.e. the plot of household relative income is insufficiently steep), and consequently household income per capita decreases in larger households. These figures are obtained for each household size by dividing column (3) in Table 1 by the share of persons in column (2), and is also plotted in Figure 1, illustrating the tendency for per capita income to fall with household size. The share of one person households and households with 8 or more persons declined sharply in Taiwan from 1966 to 1972, as shown in column (1) of Table 1. The decline in large households continues throughout the period. whereas the proportion of one-person households revives to the initial level during the 1980s, as resources became sufficient to permit more young and old individuals to maintain separate households.

Kuznets (1980) summarized income inequality in this case by the Total Disparity

Table 1

Distribution of Income among Households and Persons in Households, by Size of Households: 1966 to 1993

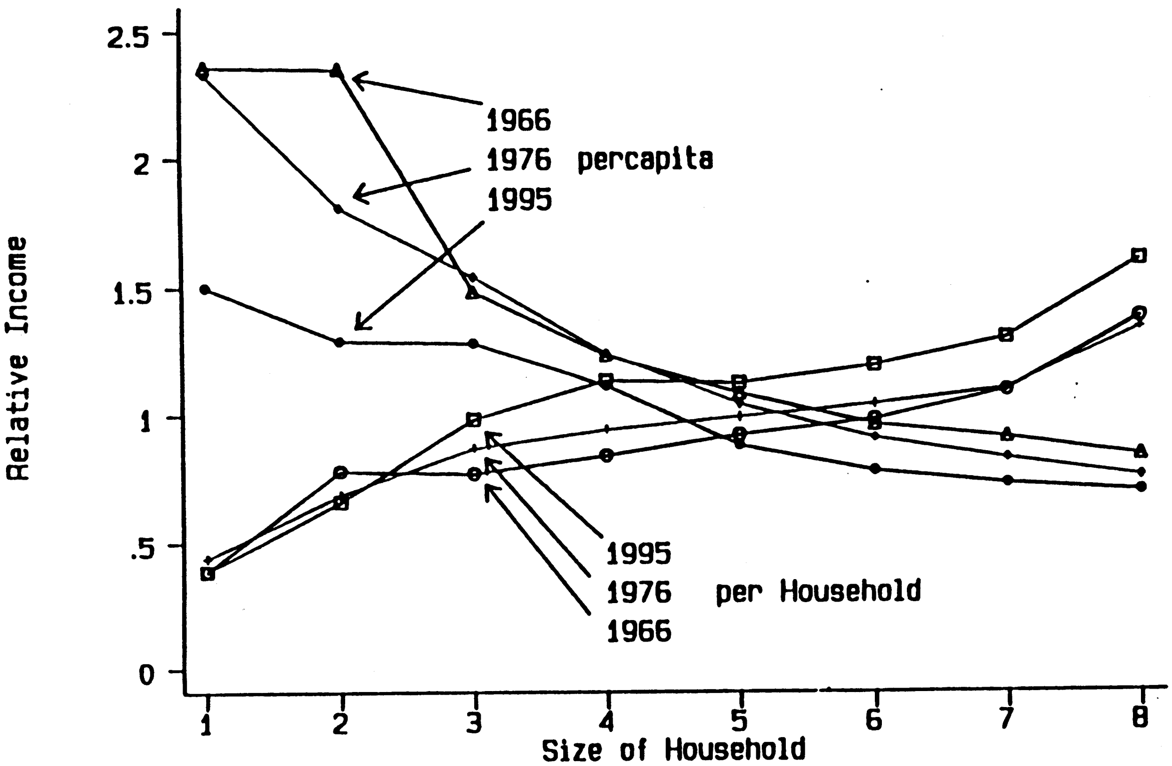
Household Size Class	1966			1972			1976			1995		
	Percent of House-holds (1)	Percent of Persons (2)	Percent of Income (3)	Percent of House-holds (1)	Percent of Persons (2)	Percent of Income (3)	Percent of House-holds (1)	Percent of Persons (2)	Percent of Income (3)	Percent of House-holds (1)	Percent of Persons (2)	Percent of Income (3)
1 person	6.6	1.1	2.6	3.3	0.6	1.4	3.2	0.6	1.4	7.8	2.0	3.0
2 persons	5.4	1.8	4.2	4.1	1.5	2.8	5.5	2.1	3.8	14.8	7.5	9.7
3 persons	7.7	4.0	5.9	9.3	5.0	7.7	9.5	5.4	8.3	16.1	12.3	15.7
4 persons	11.5	7.9	9.7	13.8	9.8	12.5	17.7	13.5	16.7	26.5	27.0	29.9
5 persons	15.3	13.0	14.0	21.2	18.9	20.9	22.9	21.8	22.6	19.5	24.9	21.9
6 persons	14.8	15.1	14.5	19.3	20.7	19.6	18.8	21.5	19.6	8.9	13.6	10.6
7 persons	14.9	17.8	16.2	12.6	15.8	13.7	10.8	14.4	11.9	4.0	7.1	5.2
8 or more Persons	23.9	39.3	32.9	16.4	27.7	21.4	11.8	20.7	15.8	2.5	5.7	4.0
Total Units[a]	2,281	13,360	96.46	2,772	15,470	167.7	9,442	49,483	1,189.6	14706	57699	1,793.5
Total Disparity Measure: Income per Household or per Person	20.8	17.2	-	12.8	19.0	-	11.9	19.7	-	20.5	19.1	-

Source: 1966 and 1972 from Kuznets (1990) Table 2.
1976 and 1995 calculated by author from Survey data files.

[a] Income total on col.(3) is in billions of current Yuan.

Figure 1

Relative Income By Household Size: Taiwan 1966, 1976, 1995



Source: Table 1.

Measure (TDM), defined as the sum of differences in percentage shares of households, and shares of income, across the distribution by household size, disregarding in the summation the signs of the differences. Kuznets compares this TDM with the Gini/Lorenz concentration ratio, because it also sums the absolute value of differences in incomes, as measured across any particular grouping of the population, e.g., by household size or by sector. From 1966 to 1972 the TDM declines markedly, as noted in the bottom row of Table 1, from 20.8 to 12.8 in terms of the distribution of income per household (col. 1), whereas the TDM increases somewhat, from 17.2 to 19.0, for the distribution of household income per capita (col. 2). Kuznets concludes "that the evidence for significant reduction in income inequality over the decade preceding 1975 does not stand up under scrutiny. The crude adjustment for size of households removes the trend suggested by the conventional distributions; and leave us with apparent constancy over the period" (Kuznets, 1980, p.264). Working from the published cross tabulations of the income and expenditure surveys from 1966 to 1975, Kuznets could appraise differences in income by household size groups, but he could not evaluate differences in income within household size groups in terms of income per capita.[2] Nor can I, until the income data are publicly available at the household level starting in 1976. Thus, it is not possible to establish conclusively whether in this early period income inequality in per capita terms decreased as did unstandardized household income inequality.

Tabulations of the subsequent micro survey data that are examined in this paper confirm roughly parallel estimates of inequality in 1976 to those Kuznets derived from published tabulations for 1972. However, the TDM across household incomes (col. 1) indicates that this unstandardized measure of income inequality has returned by 1995 to the level of 1966, or 20.5. On the other hand, the inequality in the distribution of household per capita income (col. 2) has remained relatively constant across this size grouping of households from 1972 to 1995 with the TDM fluctuating narrowly around 19. This empirical regularity for Taiwan is the main finding of this paper, and is documented more fully in the subsequent analysis. Kuznets appears to have viewed the early evidence of decline in the relatively low level of income inequality in Taiwan as based on an inadequately standardized index of income inequality. The increase in household income inequality in Taiwan from 1980 to 1995, by the same flawed measure, may also not be a satisfactory indication of change in social or economic inequality?[3] Kuznets (1980) was skeptical that the decrease in inequality in income distributed by households from 1966 to 1975 was significant, and I would argue that a similar caution should be exercised before interpreting the increase in inequality during the 1980s in household incomes as a real loss in social welfare due to inequalities in consumption opportunities.

Yet it is common practice in studies of income inequality to report the income distribution across households and not incorporate in the analysis the compositional differences in households, of which size is only the most obvious (e.g., Chiou, 1996; Chu and Jiang, 1997). In summarizing inequality in terms of per capita household income, Kuznets' TDM implicitly weighted inequality by persons in col. 3 of Table 1, rather than by households as in col. 1. This shift toward more democratic weights, although reasonable, does not preclude more refined standardizations for household composition. An intermediate position between dividing household income by size and ignoring its size might be to divide household income by the number of adults in the household, because adults might be income earners and decision makers regarding how families are constituted, in which case the number of adults in the household could be viewed as the social "weight" of the household. As argued later, research on household income distributions should ultimately account for the observed changes in family formation and household composition, because these changes are a response of individuals to labor market opportunities, the households' skills and nonhuman wealth, private transfers between altruistic or cooperating households in the extended kinship system, and public tax and transfer schemes.

3. Measures of Inequality

There are many ways to measure inequality; I focus here on only a few (Kuznets, 1989). First there is the distinction between household and individual inequality. The household includes individuals who, to some degree, pool income and share consumption. The traditional definition of a household was that its members shared a kitchen or the consumption of food in particular. Transfers across households in an extended family network can separate production from consumption, but a major function of the household is providing consumption for those residing <u>within</u> the household. Those members of the household who are less productive, either because of their stage in the life cycle, as with children and the elderly, or because they are temporarily unemployed or disabled, or more permanently incapacitated, or specializing in home production, as is a housewife, are presumably supported in their consumption requirements of goods purchased in the market by the more productive members of the household. To assess the distribution of welfare in a society, the basic unit for the study of consumption is, therefore, the household or coresidential family. Exchanges between households are not studied here, because both sides of the exchange are not observed in this or most standard household sample surveys.

The second distinction in defining inequality is the time period over which income flows are averaged. For many purposes it would be useful to measure inequality in income over a lifetime, for if financial markets can redistribute these resources over time to when they are most urgently needed for consumption and investment purposes, the lifetime constraint is most relevant for welfare comparisons. Shocks to income in one period that are offset by shocks in the opposite direction in another period might then not be construed as necessarily affecting lifetime inequality, and if these shocks are not partially anticipated, people may still be able to insure themselves against generally expected perturbations of nature. But most survey data are collected for much shorter intervals, typically one year. Annual income will be the primary basis for assessing inequality in this paper, though it is expected to overstate longer term welfare inequality. Comparing inequality in consumption to that in annual income provides one check on the possible magnitude of this upward bias.[4]

The third issue that is central to this paper is how to adjust household income for the composition of the household. The first procedure is to neglect the composition or size of households and treat all family/households as a comparable welfare unit. The second most common procedure is to divide income by the number of persons in the household, ignoring how consumption requirements vary according to the characteristics of persons or how the composition of the household may reflect the preferences of its members. More refined approximations might introduce an adult equivalent share for children that might be between about .3 and .5, and some studies even adjust for the different calorie requirements of adult males and females, presumably related to their weight or customary work (Deaton and Muellbauer, 1980; Fogel, 1994). Evidence can also be marshalled that suggests there are economies of scale in consumption which increase with household size.[5] The third procedure considers household income per adult, to measure income distribution independently of differential fertility. In this third approach the number of adults in the household becomes the population weight, and children are not attributed consumption requirements in reckoning inequality in well being.

In these three methods for dealing with household composition, it is implicitly assumed that there is an equal sharing within the household that gives all members access to the same welfare level from their consumption. A fourth issue in measuring inequality is how to measure and interpret welfare inequality within the household. Adult men and women may not have the same claim on resources within the household, and this may vary across households and overtime. The example above of lower calorie requirements for the average woman than man is based on weight-determined body metabolism rates, which may also be linked to lower physical productivity of women than men in some physically demanding (i.e., high calorie) tasks (Schultz, 1996). There could be another sharing rule less closely related to biological work capacity and more directly linked to the relative productivity or bargaining power of women and men, given their technological production opportunities and their training and skills. The educational attainments of men and women in Taiwan are becoming more similar, and this may have implications for the gender gap in wages for those who work in the labor force, as well as the gap in marginal productivity in home production.[6] The convergence in educational attainment of women and men may thus foreshadow a modification in the rules for sharing of household output and consumption.

The fifth issue is how to summarize the dispersion in the distribution of income. Two indicators of inequality or dispersion are consulted here. The first is the Gini concentration ratio (G), which is the sum of the absolute value of the differences in income (y) between all possible pairs of households, divided by twice the mean income (m) multiplied by the number of households (n) squared:

$$G \;=\; [1/(2mn^2)]\sum_{i=1}^{n}\sum_{j=1}^{n} |\, y_i - y_j \,| \;\;, \tag{1}$$

where the subscripts i and j run across all n households. The Lorenz curve provides a visual analogue for the concentration ratio and provides the intuition for why $0 \leq G \leq 1.0$.

The second indicator is the log variance (ℓv), which is the average squared deviation of a household's logarithmic income $\ell v \;-\; [1/$ from the population's (geometric) mean logarithmic income.

$$\ell v \;=\; [1/n]\sum_{i=1}^{n} \left(\ell n(y_i) \;-\; \overline{\ell n(y)}\right)^2 \;, \tag{2}$$

where $\overline{\ell n(y)}$ refers to the mean log income for all n households. This measure of dispersion assigns greater weight to equal transfers of income from the mean to the poor than from the rich to the mean, because the logarithmic transformation collapses the income scale to proportional variation. The deviations of the individual from the population mean log income are also squared in the log variance, whereas the absolute values of the deviations are arithmetically summed in the case of the Gini. This implies that the log variance also assigns greater weight than does the Gini to outliers.[7]

An attraction of the log variance as a measure of income inequality is that it can be directly decomposed into the shares of the log variance attributable to various groupings of the household that may be associated with income (Fisher, 1930). An unfortunate feature of the log variance is that families that do not have positive income in the reference period cannot be included, because the logarithm of a nonpositive number is undefined. Assigning all persons at least some minimum positive income level is an arbitrary solution to this difficulty, but not a conceptually attractive one. Over a lifetime, a negative income constraint is implausible, but for shorter periods, savings and dissaving as well as interhousehold and public transfers allow for the smoothing shocks to income as they affect consumption, and of course, a family can legitimately report negative capital income or self-employed income in any particular year, although only two households do in the data analyzed here.

Two steps in my analysis motivate the choice of log income and log variance as the measures of welfare levels and inequality. In estimating earnings functions, it is widely found that the dependent variable is fit by the standard conditioning variables better if earnings are expressed in logarithmic rather than arithmetic form (Mincer, 1974; Heckman and Polachek, 1974). The errors to this earnings regression are also more nearly normal and homoskedastic when the earnings are expressed in logarithms. Tests of statistical significance in the analysis of variance methods are also motivated by the assumption the variates are normally distributed (Fisher, 1930). In Figures 2 and 3 the actual distributions of household log income per adult are shown for 1976 and 1995 from the Taiwan Survey of Personal Income Distribution (DGBAS, 1989). The histograms of the data are compared to the frequency distribution (plotted as a solid line) implied by the parameters of the fitted lognormal distribution. The approximation is reasonably close for both the individual full-time earnings and household per adult income (shown) and lognormality cannot be rejected (D'Agostino et

Figure 2

Fraction of Households by Household Income per Adult: Taiwan, 1976

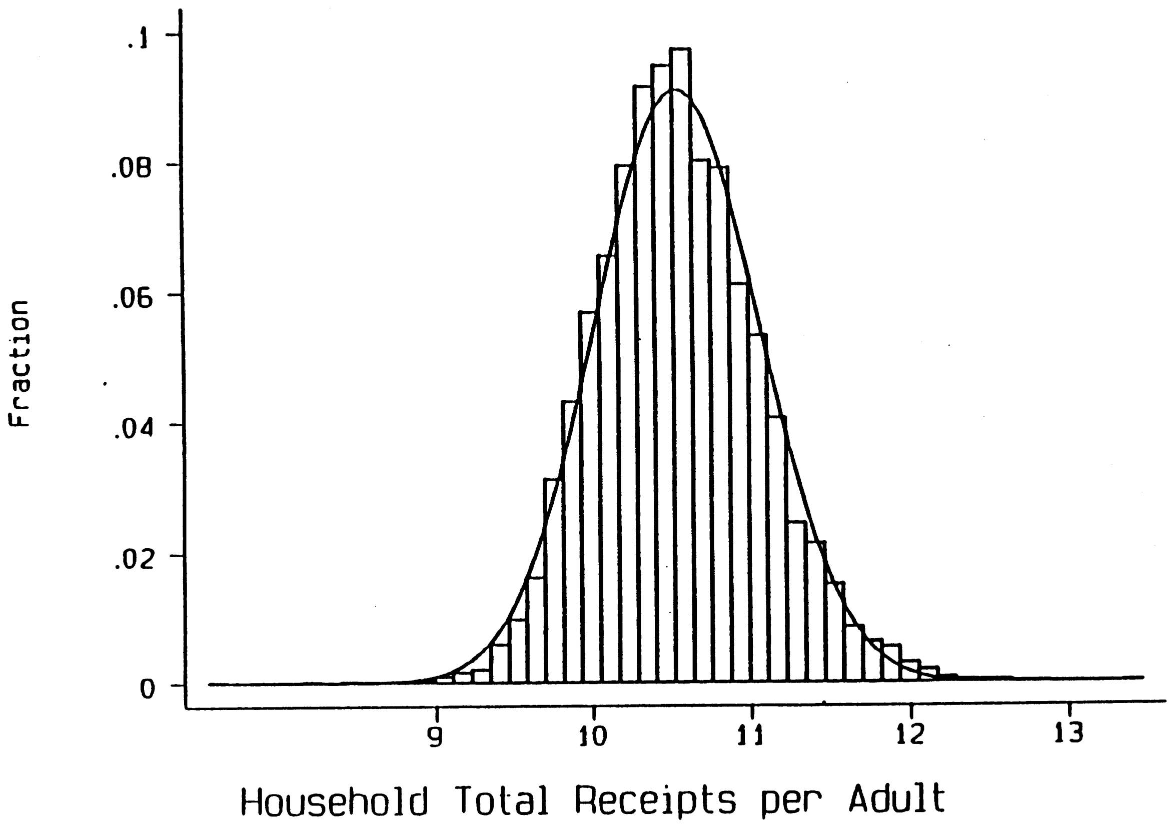
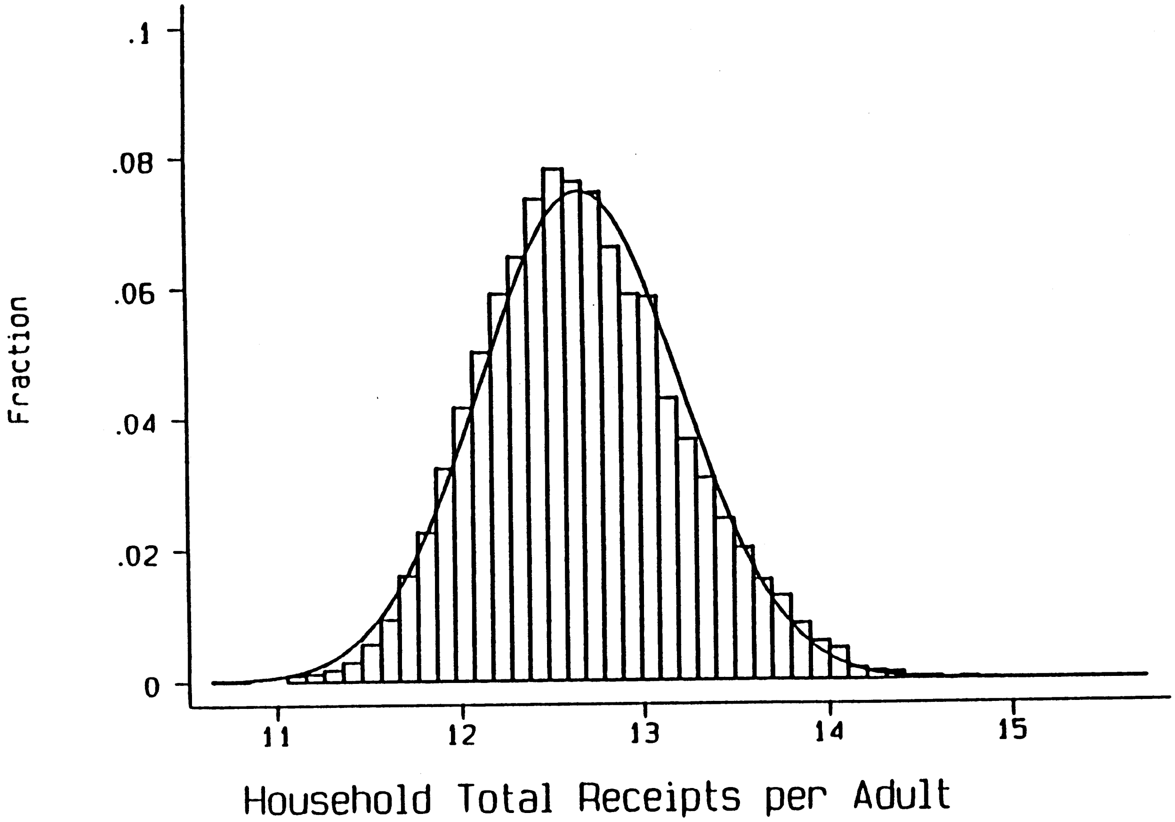


Figure 3

Fraction of Households by Household Income per Adult: Taiwan, 1995



al., 1990). If household incomes are actually distributed log normally, alternative measures of inequality — e.g., Gini, log variance, coefficient of variation, quantiles — are monotonic analytic functions of each other. In particular, the Gini can then be computed from the log variance (e.g., Aitchison and Brown, 1957, Appendix A; Deaton and Paxson, 1994). However, if measured household income deviates sufficiently from lognormal, these alternative measures of inequality could rank inequality differently and vary in opposite directions over time. This does not seem likely in the case of Taiwan if household income is standardized by either the number of adults or persons in the household and inequality is summarized by the variance in the logarithms of income.

4. Data and Trends in Taiwan

Data are drawn from the Survey of Personal Income Distribution for Taiwan (or Family Income and Expenditure Survey) conducted by the Directorate-General of Budget, Accounting and Statistics (DGBAS), Executive Yuan (DGBAS, 1989). This household random stratified survey was initiated in 1964 and expanded to include disaggregations of rural and urban areas of Taiwan from 1966. The questionnaire has changed relatively little after 1966, including for each individual in the household their income sources, economic status, industrial sector of employment, marital status (after 1987), education, sex, and age, plus detailed categories of household consumption expenditures and outlays on durable goods. Before 1976, only published cross tabulations of the survey are available, but from 1976 onward the individual household data files are available from DGBAS. The survey methodology appears to have been consistently applied from 1966 onward (Fei, Ranis and

Kuo, 1979), although the size of the survey has grown from approximately 3,000 families in 1964, to 9,500 families in 1976, to 15,000 since 1980. Approximately 50,000 individuals are enumerated in 1976, increasing to 75,000 in 1985, and decreasing to approximately 58,000 by 1995, as average household size declined. Household members are identified by their relationship to the head of household. This leaves some ambiguity in matching husbands and wives, if one of them is not the head of household. Additional information on the identity of spouses in the household is added after 1987. Consequently, to preserve the same basis for comparing (matched) couples for all years of the survey, one sample of couples used to infer full-time earnings is later restricted to couples who are also household heads.[8]

Two changes in family composition in this period are portrayed in the last two rows in the panel I of Table 2. The number of adults (i.e., persons over age 14) per family has decreased from 3.24 in 1976 to about 2.95 by 1995. This includes a marked decline in the proportion of adults residing with their elderly parents or parent-in-laws.[9] The greater frequency of extended family living arrangements noted in East Asia (Kuznets, 1989) appears to be diminishing, at least in Taiwan after 1976.

The other change in family composition is the decrease in number of children per adult from .62 to .32 from 1976 to 1995, which is more striking when only prime age adults 15 to 65 are in the denominator of the ratio, who are more likely to be the parents of the household's children. The secular evolution of the demographic transition in Taiwan is most readily measured in the decline in birth rates that followed the abrupt reduction in mortality after the Second World War. The crude birth rate peaked at 50 per thousand persons in 1951, and declined to 38 by 1961, to 26 by 1971, and then slowed its descent to 23 in 1981, and

Table 2

Households, Incomes and Inequality, with Three Alternative Adjustments for Household Composition

	1976	1980	1985	1990	1993	1995
I. Composition of Families						
1. Number of Families	9437	14697	16430	16434	16434	14706
2. Number of Adults (Age 15+)	30545	46307	51548	49003	49687	43409
3. Number of Persons	49483	71231	75496	68846	67227	57699
Adults per Family (2/1)	3.24	3.15	3.14	2.98	3.02	2.95
Children per Adult ((3-2)/2)	.62	.54	.46	.40	.35	.32
II. Income Level:						
4. Mean Income Families (in thousands of current NT)	126.	254.	362	593	836	996
5. Mean Income Families per Adult (in thousands of current NT)	38.8	80.7	115	199	277	337
6. Mean Income Families per Capita (in thousands of current NT)	24.0	52.5	78.1	142	204	254
7. Mean Log Income Families	11.601	12.308	12.645	13.118	13.457	13.635
8. Mean Log Income Family per Adult	10.433	11.159	11.511	12.052	12.381	12.587
9. Mean Log Income Family per Capita	9.943	10.724	11.126	11.711	12.086	12.310
III. Income Inequality						
10. Gini Coefficient Families	.2892	.2845	.2977	.3134	.3150	.3131
11. Gini Coefficient Family per Adult	.2872	.2795	.2974	.3026	.3023	.2965
12. Gini Coefficient Family per Capita	.2947	.2973	.3015	.3023	.2959	.2887
13. Variance of Log Families	.2771	.2915	.3177	.3752	.3969	.3838
14. Variance of Log Family per Adult	.2488	.2637	.2698	.2794	.2829	.2634
15. Variance of Log Family per Capita	.2586	.2676	.2758	.2793	.2678	.2491
16. Consumer Price Index (1991=100)	48.3	71.5	86.6	96.5	107.5	116.0

Source: Rows 1-15 calculated by the author from Survey data (DGBAS, 1989).
Row 16: DGBAS Statistical Yearbook (1994).

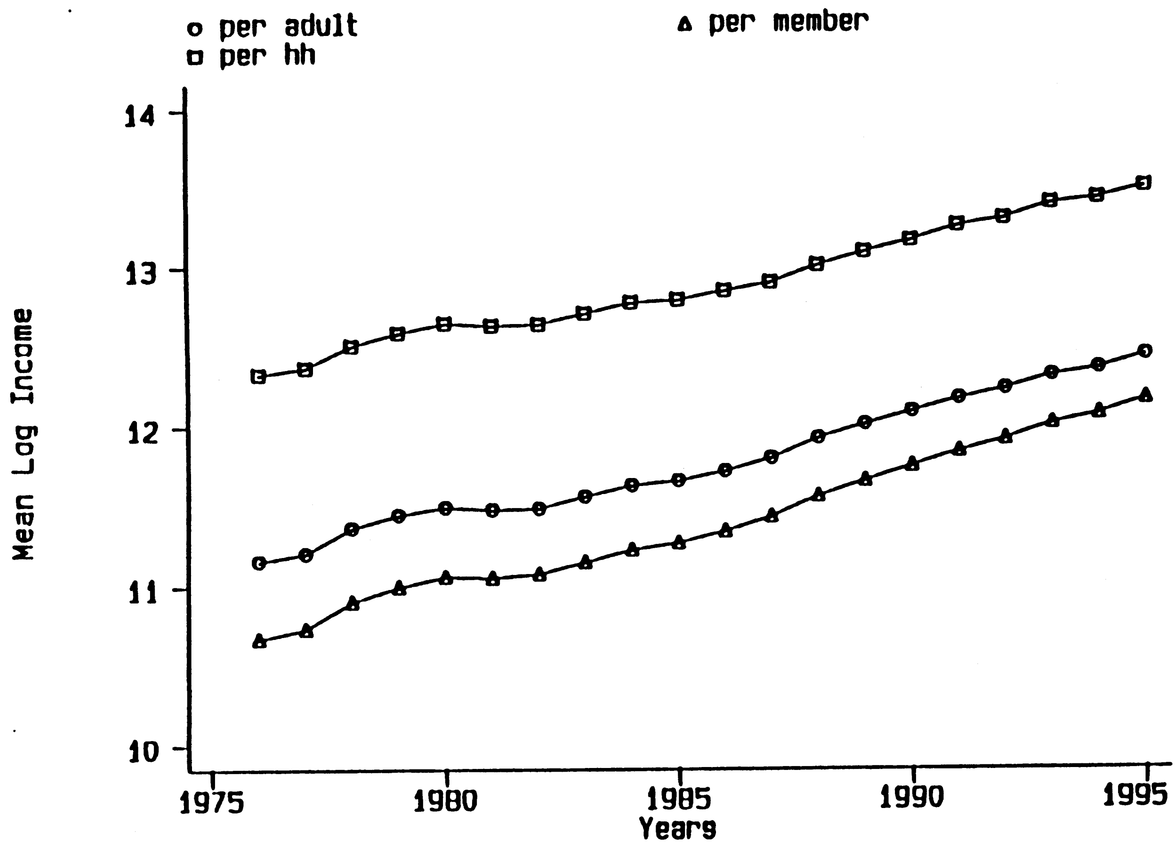
stabilized at about 16 after 1991. Primarily as a consequence of this decline in crude birth rates the proportion of the population under 15 years of age fell from 45 percent in 1965 to 25 percent in 1993. Expected lifetime fertility for women, or the total fertility rate defined as the sum of current age-specific birth rates from age 15 to 45, fell from 5.1 children per woman in 1964 to less than replacement of 1.8 children after 1986. The corresponding percent of the population elderly, or age 65 and over, increased from 2.6 percent in 1965, to 4.3 percent in 1980, to 7.1 percent in 1993, and is projected to continue to increase rapidly in the future (DGBAS, 1994; Taiwan, Ministry of Interior, 1991).

Household income per capita reported in panel II of Table 2 has increased eight fold in the 17 years from 1976 to 1993, whereas income per adult increased more slowly by seven-fold, due to the declining proportion of children. The consumer price index increased 122 percent, reducing the real income growth per adult to 320 percent, suggesting that real income per adult has grown at a continuously compounded rate of 7.1 percent per year. Figure 4 plots the annual estimates of the mean log real incomes in 1991 NTs, based on the log of income per household, household income per adult, and household income per member.

The family would seem the appropriate unit to measure welfare inequality, but the formation of families is itself endogenous to income opportunities, to some degree. They can form and subdivide to realize benefits for their members, whether these benefits are observed in income receipts, or are related to economies of scale in consumption or production, or psychic benefits of privacy, (dis)economy of scale in some forms of consumption, or a specific matching of individual demands for public goods, such as number and quality of children.

Figure 4

**Time Trend in the Log Real Income:
Per Household, Per Adult, and Per Member**



Income per adult normalizes household resources prior to any fertility decisions. It treats as equally well-off two couples that have the same income, regardless of whether they have decided to have, say, one or two children. This approach ignores the extent to which couples have more or less children than they want, for the welfare of couples who do not obtain their desired fertility should be lower than other observationally equivalent couples, i.e., with the same income and fertility. Society may also believe that children should receive a minimum material standard of living. Measures of social inequality based only on income per adult could thereby neglect the extent to which children are disproportionately in poorer families ranked by income per adult.[10] By the same logic, society may decide the consumption needs of the elderly should be treated in a different manner from other "prime aged" adults. But since savings during productive adult years is a means for sustaining consumption in retirement years, the treatment of elderly with other adults, regardless of their current participation in income earning activity, is arguably appropriate. Family income may also be divided by the number of persons in the family, including children as equal claimants to adults on consumption and savings. The per capita and per adult measures of family welfare bound my measures of inequality, and presumably if intermediate adult "equivalent" weights were defined for children between zero and one, the resulting measure of equivalent inequality would be somewhere between the two reported here.

Differential fertility by the level of adult incomes may also be a factor affecting per capita income inequality over time. The demographic transition in low-income countries is associated with a period of rapid population growth that could have increased income inequality, if fertility decreased faster among the rich than the poor, concentrating investments in human capital per child among the upper classes of subsequent generations (Schultz, 1971). Although the survey data analyzed here are not well-designed to measure differential fertility or child investments, some insight into these behavioral questions can be gleaned from a simple analysis of the distribution of children by household income per adult.

5. Changes in Household Income Inequality

The third panel in Table 2 compares the two summary indexes of inequality, using the three measures of economic welfare in the family described above. All annual estimates of the Gini coefficient are plotted in Figure 5 according to the three household bases of welfare. The Gini coefficient for entire households has increased 11 percent from its low in 1980 of .285 to a peak in 1993 of .315, before stabilizing. The Gini in terms of household income per adult has increased 8.3 percent from 1980 to 1990, and has fallen since 1991. If income per capita in the household is the relevant measure of welfare, the Gini has been nearly constant from 1978 to 1991, when it starts to decline.

Figure 6 plots the annual estimates of the variance of log incomes according to the three methods for normalizing for the composition of the household. The variance of log of total household income increases 39 percent from 1976 to 1995. Based on household income per adult, the increase is 5.9 percent, and on an income per capita basis the variance of logs decreases 4 percent over the entire period. If we compare the inequality based on household income per adult and per member (adding children), our measure of inequality is higher for per capita inequality in the early years, but after 1987-1990 the addition of children to the welfare comparisons reduces the variance of log income slightly. Conversely, inequality is

Figure 5

Time Trend in Gini Concentration Ratio of Income: Per Household, Per Adult, and Per Member

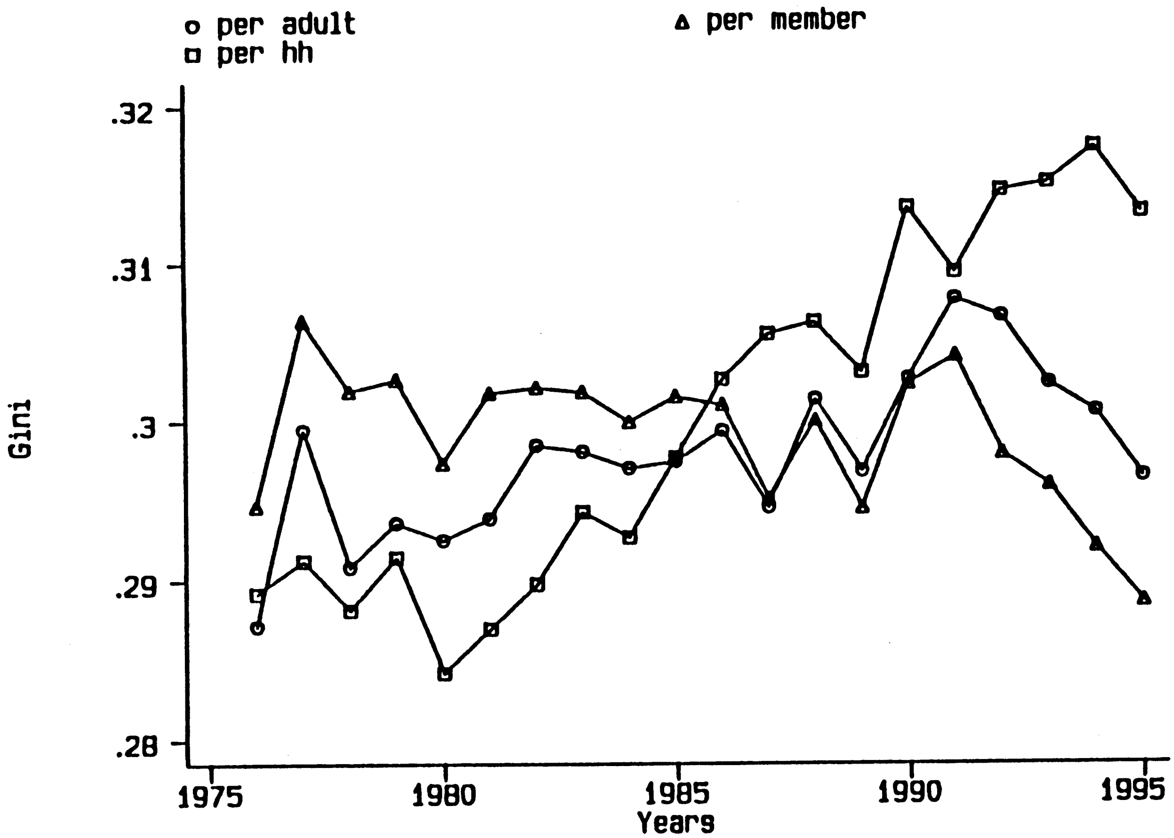
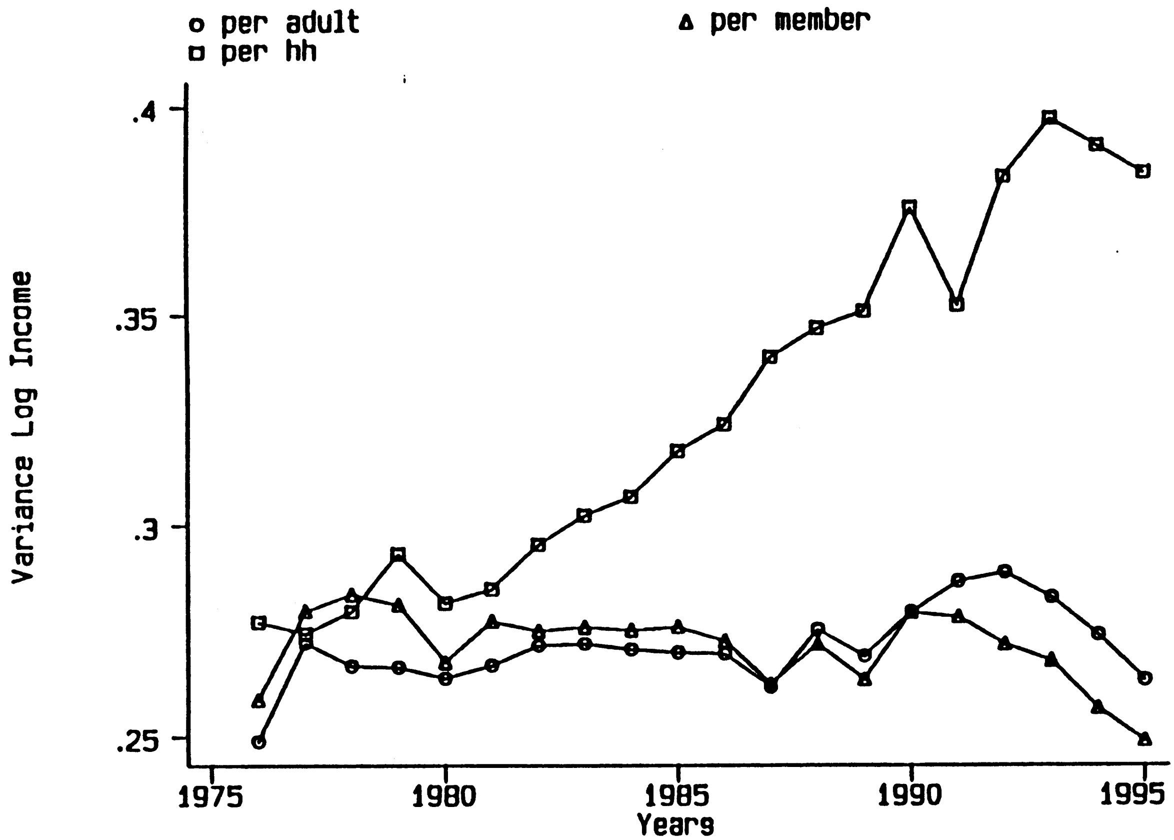


Figure 6

Time Trend on Log Variance of Income: Per Household, Per Adult, and Per Member



smaller when family welfare is represented by either income per adult or per person rather than simply as household income. The important distinction is that when household size, measured either as number of adults or as number of persons, is used to normalize household income, the sharp rise in household inequality is essentially eliminated. <u>The apparent increase in the variance (or Gini) of log household incomes in this period appears to be associated with changes in the composition of households</u> and does not reflect an unambiguous change in the distribution of economic welfare, just as Kuznets argued in the earlier period of 1964-1975.

Income Effects on the Demand for Household Composition

If individuals modify the composition of the households within which they live as their income opportunities, prices, and technologies change, it may be informative to describe how household composition is related to the income level of adults. The dynamic processes of marriage, fertility, and extension of families to include adult relatives is exceedingly complex and is not expected to respond uniformly to income opportunities. Moreover, these lifecycle changes in family composition are expected to be more closely related to long-term income opportunities than to annual income. To better proxy this concept of permanent income, the level of household expenditures per adult is examined in addition to annual income per adult. Three age groups in the family are distinguished: A or the number of prime aged adults (age 15 to 65), C or the number of children (age less than 15), and E or the number of elderly (over age 65). To permit an additive decomposition of the effects of income on log of family size, three components are defined that sum to the total: ln A; ln((C+A)/A); and ln(E+C+A)/(C+A). The arithmetic forms of these family components (A, C/A, and E/A) are examined as alternative dependent variables, because they are more intuitive but their effects on log family size are less readily decomposed. The regression coefficients on household expenditures (income) per adult are reported in Appendix Table A-1, for 1976 and 1995, where controls are included for the age and age squared of the household head (defined as the person with the most income in the household). The samples are restricted to households for whom the head is between age 30 and 50. This restriction on age is designed to make C/A more closely approximate lifetime fertility, and E/A more reasonably represent the extension of the adult children's household to accommodate elderly relatives who are predominantly parents (cf. Table A-2).[11]

The association between the various household components and expenditures or income is not uniform; it is negative with respect to (prime) adult size of household and extension of the household to include elderly, but positive with respect to fertility. The relationship is as expected generally more statistically significant when income is proxied by expenditures per adult, and these are consequently my preferred estimates. The elasticity of adult size with respect to expenditures per adult decreases in absolute value from -.40 to -.32 from 1976 and 1995. The elasticity of elderly extension (approximated by E/A) decreases in absolute value from -1.48 to -1.16 during these two decades when the number of elderly per prime aged adult in these households more than doubled, from .046 to .106. Finally, the elasticity of fertility (approximated by C/A) increases from .45 to .51 in a period when this measure of fertility declined by 40 percent.

The negative income elasticity with respect to adult size of household could be due to a decrease in the share of welfare generated by market production as single men and women are married and set up their own households, at which time women increase their allocation of time to home production and child care. But the negative association of adult size of household and earnings (income) per adult also exists for households with two or more prime aged adults. Economies of scale in production and consumption could provide these larger families with more welfare from the same expenditures. The adult size and elderly inverse relationships with income could also be interpreted as a demand for privacy, viewed as a normal good, that is sacrificed when the number of adults living in a household increases. Different preferences for expenditures between younger and older generations could also grow more pronounced with rising incomes, suggesting another reason why the elderly increasingly maintain their own households as development progresses (Hayashi, 1995).

The positive relationship between fertility and household expenditures per adult is unexpected in this post-demographic transition society where fertility levels are slightly below population replacement levels. Children appear to be a superior good, and households (and couples) with twenty percent higher expenditures (income) per adult have on average 8 to 10 percent more children.

6. Changing Age Composition of Households and Inequality

In the wake of the demographic transition when mortality and fertility secularly decrease, the proportion of the population in older age groups increase. This pattern would be reinforced in comparisons of household income by the growing propensity of the elderly to maintain separately enumerated households. Because it is widely observed that earnings and income inequality tend to increase with age (Mincer, 1974), these developments in Taiwan might have contributed to increasing measured income inequality. A simple method to assess how the changing age distribution of households may have affected measured income inequality is to decompose the variance of logs of incomes per adult across all households into within age-group variances and between age-group variance components associated with the lifecycle income profile (Schultz, 1965). It is then straightforward to evaluate the contribution of the changing age weights over time to measured inequality, assuming that the within age group inequality and income age profiles are unchanged. Of course it is also possible that the changing age composition of the population affects both the age profile of incomes and within age-group inequality. These latter possibilities are not considered here.

Table 3 performs this decomposition of the variance in log incomes per adult in the first and last year of my data, 1976 and 1995. The first column reports the percentage of all households in each age group. The next two columns report the mean of log incomes and variance of log incomes within the age group, where the mean log income and variance of log income for the sum of all ages is identical to that in Table 2. Column (4) shows the share of the overall log variance that is accounted for by the within age cohort log variance weighted by the population share, and column (5) the share due to the between age cohort and overall mean log income, squared and weighted by the population share (see Table notes). The sum of the figures in columns (4) and (5) add up to the log variance for all ages combined, as reported at the top of column (3).[12]

Several common regularities in income inequality across age groups can be seen from Table 3. In 1976 one observes relative inequality in income increases with age from .16 for those age 15-19, to a peak .26 for ages 35-39, declining then before it rises again to .30 after

Table 3

Decomposition of Log Variance of Household Incomes per Adult by Age Weighted by Number of Adults in Household

	1976					1995				
Age of Household Head*	Percent of Adult Population	Mean Log Income	Variance of Log Income	Contribution of Group to Variance of Log Income		Percent of Adult Population	Mean Log Income	Variance of Log Income	Contribution of Group to Variance of Log Income	
				Within	Between				Within	Between
	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
All Ages	100.00	10.433	.2488	-	-	100.00	12.587	.2634	-	-
15-19	0.62	10.103	.1568	.0010	.0007	0.34	12.089	.1708	.0006	.0008
20-24	3.28	10.231	.1861	.0061	.0013	2.99	12.348	.1491	.0045	.0017
25-29	9.54	10.488	.2339	.0223	.0003	7.66	12.561	.1519	.0116	.0001
30-34	10.40	10.650	.2530	.0263	.0049	12.82	12.714	.2148	.0276	.0020
35-39	13.01	10.616	.2613	.0340	.0044	15.34	12.790	.2704	.0415	.0063
40-44	16.18	10.402	.2288	.0370	.0002	17.69	12.589	.3001	.0531	.0000
45-49	18.84	10.355	.2073	.0391	.0011	14.93	12.486	.2463	.0368	.0015
50-54	13.07	10.363	.2115	.0276	.0006	9.63	12.565	.2179	.0210	.0000
55-59	8.28	10.365	.2412	.0200	.0004	7.31	12.610	.2320	.0170	.0000
60-99	6.77	10.303	.3007	.0204	.0011	11.28	12.400	.2953	.0333	.0040

Notes: Column (4) = Col.(3)* Col.(1)/100.
Column (5) = [(Col.(2)-Col.(2 for all ages))**2]*Col.(1)/100.

*Household head is defined as the individual who has the largest annual income.

retirement for those over age 59 and remain a household head. This lifecycle pattern in within age-cohort inequality has become more equal by 1995 through age 30-34, but thereafter it has become slightly less equal from age 35-39 to age 50-54. The mean log income also increases until age 35-39 in both 1976 and 1995, and then declines slowly with age, until around retirement. Because the mean income profile is relatively flat across ages, increasing by only about 40 percent from age 20-24 to age 35-39 and back to retirement, the between age-group component contributes a relatively small share to the overall log variance (col. 5). The within age-group variance component (col. 4) grows more substantial for the middle aged and elderly. As the age composition of the population in Taiwan has shifted toward the older ages, the log variance of income may have tended to increase.

To quantify this effect of the changing age composition of households related to the demographic transition, assume that the within and between age-cohort inequality components from 1995 were weighted by the 1976 age distribution of households (col. 1). Then the resulting 1976 counterfactual log variance would have been .2537 compared with the actual 1995 value of .2634. Alternatively, if the cohort inequality components from 1976 are weighted by the 1995 age distribution, the resulting counterfactual log variance for 1995 would have been .2566 compared with the actual 1976 value of .2488. Thus, of the actual increase of 5.9 percent in the variance in log incomes per adult across all ages from 1976 to 1995 from .2488 to .2634, 2.0 percentage points of the increase in overall log variance would have occurred due solely to the change in age distribution, holding constant the within and between cohort components of inequality at their initial levels. Conversely, 2.7 percentage points of the increase could be attributed to the age composition change evaluated at the final inequality levels. One-third to one-half of the increase in log variance of income per adult across households in Taiwan in this period is thus attributable to the changing age composition associated primarily with the demographic transition. The increases in inequality due to the age groups 35 to 44 and 60+ outweigh the inequality decreases in the intermediate ages. Because Taiwan absorbed an unusually large immigration of adults in the 1940s and 1950s from the Chinese mainland, the nearly doubling of the proportion of households over age 60 from 1976 to 1993 is larger than might be expected in a more typical, closed population that is experiencing a rapid demographic transition.

7. Changing Time Allocations of Women and Household Full Income

The share of women age 15-65 who are in the labor force in Taiwan increased from 37.6 percent in 1976 to 44.9 percent in 1993 (DGBAS, 1994, Table 27) and the share of women working as full-time employees doubled from 1976 to 1995 (cf. later Tables 6 and 7). This increased participation of women in the labor force may be associated with the decline in fertility and a reduction in the time mothers allocate to child care. Because my data on market income assign no value to the leisure activities of adults or the productivity of their time outside of the labor force, the increased participation of women in the labor force might have occurred mostly among lower income groups and thus contributed to a more equal distribution of market income while adding to the inequality in nonmarket production.

To quantify the impact of women's increased labor force participation on the distribution of household welfare, I estimate in this section the shadow value of the household's resources, before the members of these households make their decisions on market labor supply. My objective is to approximate the household's "full income" endowment (Becker, 1965). To implement this approach I must infer the shadow wage for all men and women, and then attribute the value of their time to all adults in the sample, regardless of their actual work decisions, and then add nonearned income to these full-time shadow wages to infer what the "full income" of the household might have been if all adults had worked in the labor force full time. For those who are primarily a recipient of entrepreneurial income that exceeds their estimated shadow full-time wage (only a few percent), they are assumed to retain their larger entrepreneurial income and do not receive their estimated shadow wage, because a part of their entrepreneurial income may be a return on capital and risktaking. For individuals over age 64, actual earnings are retained and no imputations of full-time earnings is attempted, because a significant share of these individuals may not be capable of working full time, due to exogenous health limitations. Despite the weakness of such an exercise that simulates a counterfactual outcome, it may provide some insights into how the changing allocation of women's time affects the distribution of market income and ultimately the distribution of economic welfare.

First, the logarithm of earnings is estimated for all full-time wage earners (those who receive their largest source of labor income from full-time employee compensation) between the ages of 15 and 64 (Table A-3, cols. 1 and 2). This sample is used because time allocated to work is not explicitly reported in the survey to calculate a wage rate, and the full-time employees are the majority of the labor force. The logarithm of earnings is estimated conditional on years of education, years of postschooling potential experience, and experience squared, separately for men and women.[13] Because these Mincer (1974) earnings functions explain only about a third of the variance in log wages, the predicted earnings has a log variance that is only a third the size of the full-time earnings in the estimation sample. A random error is therefore added to each individual's predicted wage that is drawn from a normally distributed random variate with zero mean and variance equal to the actual sample log variance of earnings multiplied by one minus the R^2 in the estimated earnings equation. According to this procedure, the variance in the simulated log wage is approximately equal to the variance of the actual earnings in the estimation sample. This method is then implemented to simulate full-time earnings for all prime-aged adults in the sample, except for entrepreneurs and the elderly as noted. After summing the full-time earnings of all adults age 15-64 in the household, each household is allocated the actual household property, transfer, and other income receipts and elderly earnings, to obtain the simulated household's full-income.

A positive correlation between the earnings potential of husbands and wives increases the inequality in full income across households, and might not be fully reflected in my initial simulation of full incomes (Kremer, 1997). Marital partners may be matched both on observable variables (e.g., education and post school experience) that already enters into the prediction of wages, and also matched on unobserved traits that could influence their shadow wage. Consequently, in a second set of estimates of full income the earnings estimation sample is restricted to married couples who are both full-time wage earners (Table A-3, cols. 3 and 4).[14] It is possible then to calculate for this sample the correlation between the log full-time earnings between spouses, which is .59 in 1976 and .56 in 1995. Education alone is correlated across all spouses even more strongly, at .66 in 1976 and has increased to .75 by 1995.[15] My predicted wages based on observables are correlated across spouses at .74 in 1976 and .75 in 1995. Finally, as expected, the residuals in the two spouses' earnings equations (i.e., unobserved factors) are also significantly correlated across couples at .27 in 1976 and .25 in 1995. The shadow wage estimates based on these full-time earner couples are therefore attributed two random errors: the first is a couple-specific (shared) error that corresponds to the unobservable productive traits reflected in the cross-spouse residual correlation, and the second is a random error that is assumed independent across all individuals. It is then necessary to scale down the size of the second iid wage error attributed to each matched spouse who has already been imputed a couple-specific error.

To incorporate in this setup the couple-specific covariance in unobservables, two assumptions are being maintained. The sample of couples who are both full-time wage earners is a representative sample of the universe of all adult couples in terms of their shadow wages, in order to avoid a sample selection bias in estimating for all person's a shadow earnings. The wage structure estimates are in fact similar for all full-time earners and matched couple full-time earners, as reported in Appendix Table A-3. The covariance of couple earnings for the joint full-time earning couples is also assumed to be an unbiased estimate of this couple-specific wage heterogeneity for all matched working and nonworking couples.

To assess the robustness of results to these potentially restrictive assumptions, I report estimates of full income in Tables 4 and 5 based on both sets of working assumptions: (1) estimates based on the earnings function for all full-time individual men and women workers by sex and ignores the differences between married and unmarried individuals in their shadow wages; and (2) estimates based on the earnings functions for married couples who are both

Table 4

Income Inequality Across Households by Age of Head, According to Actual Income Receipts and Two Simulations of Full Income: 1976

Index of Inequality and Concept of Household Income	Age Group										
	All Persons over Age 14	15-19	20-24	25-29	30-34	35-39	40-44	45-49	50-54	55-59	60+
1. Number/Proportions											
Households	9,441	.006	.030	.106	.133	.162	.154	.163	.110	.071	.066
Adults	30,557	.006	.033	.095	.104	.130	.162	.188	.131	.083	.068
Persons	49,501	.005	.026	.089	.130	.176	.171	.171	.109	.067	.055
2. Gini Ratio/Household											
Actual Income	.289	.222	.256	.271	.276	.261	.259	.267	.304	.338	.421
Simulated/Individual	.300	.286	.311	.313	.287	.263	.251	.267	.289	.315	.394
Simulated/Couple	.322	.286	.313	.320	.310	.283	.279	.298	.321	.319	.400
3. Gini Ratio/Adult											
Actual Income	.287	.216	.247	.274	.285	.286	.274	.261	.262	.287	.338
Simulated/Individual	.224	.211	.193	.215	.226	.243	.210	.214	.209	.225	.259
Simulated/Couple	.213	.209	.187	.212	.239	.237	.204	.196	.188	.204	.244
4. Gini Ratio/Person											
Actual Income	.295	.203	.262	.296	.309	.282	.268	.270	.282	.324	.374
Simulated/Individual	.283	.259	.254	.283	.295	.271	.247	.258	.265	.274	.314
Simulated/Couple	.297	.224	.256	.291	.314	.280	.270	.272	.266	.271	.303
5. Log Variance/Household											
Actual Income	.277	.182	.209	.229	.223	.211	.221	.236	.328	.416	.582
Simulated/Individual	.336	.277	.337	.314	.251	.232	.219	.256	.349	.439	.806
Simulated/Couple	.382	.323	.345	.326	.296	.262	.283	.324	.442	.459	.816

Table 4 cont.

	Age Group										
Index of Inequality and Concept of Household Income	All Persons over Age 14	15-19	20-24	25-29	30-34	35-39	40-44	45-49	50-54	55-59	60+
6. Log Variance/Adult											
Actual Income	.249	.157	.186	.224	.253	.261	.229	.207	.212	.241	.301
Simulated/Individual	.161	.147	.116	.146	.160	.189	.133	.138	.141	.158	.276
Simulated/Couple	.151	.161	.112	.146	.180	.177	.131	.122	.122	.134	.263
7. Log Variance/Person											
Actual Income	.259	.122	.208	.270	.279	.243	.221	.221	.246	.307	.365
Simulated/Individual	.266	.221	.215	.270	.277	.238	.204	.216	.236	.232	.343
Simulated/Couple	.305	.162	.232	.288	.318	.259	.253	.256	.257	.250	.330

Table 5

Income Inequality Across Households by Age of Head, According to Actual Income Receipts and Two Simulations of Full Income: 1995

	Age Group										
Index of Inequality and Concept of Household Income	All Persons over Age 14	15-19	20-24	25-29	30-34	35-39	40-44	45-49	50-54	55-59	60+
1. Number/Proportions											
Households	14,706	.003	.025	.072	.136	.179	.170	.118	.077	.066	.153
Adults	43,409	.003	.030	.077	.128	.153	.177	.149	.096	.073	.113
Persons	57,699	.003	.025	.071	.147	.200	.192	.129	.079	.060	.093
2. Gini Ratio/Household											
Actual Income	.313	.231	.256	.235	.258	.262	.271	.277	.306	.352	.438
Simulated/Individual	.288	.224	.266	.269	.258	.228	.223	.229	.250	.298	.440
Simulated/Couple	.294	.221	.267	.262	.260	.240	.232	.234	.260	.306	.441
3. Gini Ratio/Adult											
Actual Income	.297	.229	.221	.222	.264	.296	.322	.294	.268	.272	.316
Simulated/Individual	.220	.154	.165	.167	.194	.217	.225	.209	.194	.203	.303
Simulated/Couple	.221	.168	.156	.174	.200	.228	.226	.204	.192	.201	.294
4. Gini Ratio/Person											
Actual Income	.289	.246	.233	.240	.278	.290	.289	.285	.280	.291	.319
Simulated/Individual	.245	.202	.193	.225	.240	.235	.227	.215	.212	.227	.306
Simulated/Couple	.248	.207	.183	.223	.248	.249	.231	.212	.214	.226	.295
5. Log Variance/Household											
Actual Income	.384	.187	.233	.184	.218	.223	.243	.270	.358	.509	.582
Simulated/Individual	.386	.243	.286	.261	.221	.173	.185	.219	.257	.363	.760
Simulated/Couple	.396	.201	.278	.252	.224	.193	.203	.239	.284	.396	.737

Table 5 cont.

Index of Inequality and Concept of Household Income	Age Group										
	All Persons over Age 14	15-19	20-24	25-29	30-34	35-39	40-44	45-49	50-54	55-59	60+
6. Log Variance/Adult											
Actual Income	.263	.171	.149	.152	.215	.270	.300	.246	.218	.232	.295
Simulated/Individual	.168	.079	.083	.094	.124	.147	.151	.129	.115	.128	.384
Simulated/Couple	.168	.093	.078	.099	.131	.162	.152	.126	.112	.129	.359
7. Log Variance/Person											
Actual Income	.249	.193	.166	.177	.236	.251	.244	.240	.239	.260	.298
Simulated/Individual	.198	.127	.123	.174	.180	.174	.166	.147	.143	.159	.363
Simulated/Couple	.206	.148	.112	.170	.195	.196	.174	.147	.148	.162	.338

full-time earners and includes the covariance in unobservables for the wages simulated for all matched couples, but includes the wage imputations to singles as before.

The simulated-individual Gini coefficients for "full incomes" of households are somewhat larger in 1976 than the actual income Gini coefficients, .300 versus .289, but become more equal by 1995 than the actual incomes of households, .288 versus .313. The same is true for the variances of log income per household, for which the inequality of full income grew more unequal from 1976 to 1995. Using the preferred household income per adult as the welfare indicator, the full or actual income Gini inequality is essentially unchanged starting in 1976 at .224 for full income versus .287 for the actual income, and reaching by 1995 .220 for full income versus .297 for the actual income. In 1976 the variance of log income per adult is much lower for full income than for actual income, .161 versus .249, and both have increased by 1995, to .168 and .263, respectively. Families with higher income per adult tend to have more full-time market workers, suggesting that lower income groups have relatively more time for nonmarket activities, such as childrearing and leisure. On a per capita basis the full income individual Gini ratio is slightly less than the Gini of actual income, .283 and .295, whereas in terms of the log variance of per capita income, the full income is slightly more unequal than the actual in 1976. But the secular trend downward is stronger for the log variance in per capita full income than for the actual market income receipts (i.e. .266 declines to .198 for full income and .259 to .249 for actual income).

Inequality increases only slightly in 1976 for household full income and per capita income inequality when the working couples are used to estimate the imputing wage equation for couples in household or per capita income comparisons, rather than rely only on all individual full-time wage earners. However, with the covariance in heterogeneity added for all matched couples, the log variance per adult income declines, and is now 40 percent smaller than for the actual income, .151 versus .249 in 1976, and .168 versus .263 in 1995. The Gini changes less. It would appear that nonmarket time is more available for lower-income households than for higher-income households. This empirical regularity is consistent with a tendency for the market labor supply of individuals and households to be typically an increasing function of their shadow wage. In other words, the own-wage effects on labor supply (not compensated for income) tend to be positive in Taiwan during this period. Inequality in full income or economic opportunities per adult, my preferred measure of inequality, including the matching of spouses on unobservables, indicates that the Gini has increased slightly from .213 to .221 in the twenty year period 1976 to 1995, and the log variance of full incomes per adult has increased from .151 to .168. Although these time trends are probably not significant in a statistical sense, the levels of inequality are distinctly lower for full income per adult by 1995 than for actual income per adult, and according to the variance in log full income per adult, inequality is already lower than market income inequality in 1976.

The comparison of inequality estimates based on full-income and actual market income suggests that individuals with higher wage earning opportunities are more likely to work in the market labor force and work more hours in this capacity. This tendency has become more pronounced over this twenty year period. The Personal Income Surveys in Taiwan do not ask questions about the numbers of hours worked, but do distinguish between working as a full-time employee, part-time employee, entrepreneur, or unpaid family worker.

Linear probability functions are estimated to discriminate into which of these four categories an individual falls (or which job source of earnings is largest), with the residual fifth category being individuals who had no reported labor market attachment, conditional on their predicted individual based full-time log earnings, as used above to impute the first measure of full income.[16] Tables 6 and 7 report these employment type regression coefficients on this log earnings variable, and the mean participation rates in 1976 and 1995 for the four job-type categories, as well as the earnings coefficient from regressions for the sum of all four work activities, called "All Working". I disaggregate the samples between single and married men and women age 15 to 64, given the empirical tendency for labor supply responses to be larger for married women than for the three other demographic groups.

In 1976 men are much more likely to be full-time employees if their market wage opportunities are greater, although even for men the All Working effect is a sum of a negative effect on working as an entrepreneur (mostly small-scale, low-paid businesses) or unpaid family worker, and a much larger positive effect on full-time and even part-time employee categories. For single women a similar pattern emerges, but for married women the net effects of improved market wage opportunities is to reduce family and entrepreneurial employment (self employed family businesses) more than it is to increase work as an employee. But by 1995, married women are responding on balance to an increase in their market wage opportunities by working much more as a full-time employee. It may be noted that while only one in six married women age 15-64 were full-time employees in 1976, the fraction had doubled to one in three by 1995, while married women working in family unpaid jobs had proportionately declined. In 1976 an increase in women's opportunity wages, due in

Table 6

Labor Force Participation Response to Own Wages of Women and Men by Type of Employment and by Marital Status: 1976

	All Persons		Single		Married	
	Wage Coefficient[a]	Mean Participation Rate	Wage Coefficient[a]	Mean Participation Rate	Wage Coefficient[a]	Mean Participation Rate
Women Age 15-64						
Full-Time Employees	.239	.259	.285	.409	.183	.167
	(19.2)		(12.0)		(13.5)	
Part-Time Employees	.0031	.002	.0006	.003	.0039	.002
	(2.18)		(.20)		(2.44)	
Entrepreneurs	-.0268	.049	-.0295	.032	-.0269	.059
	(4.16)		(3.18)		(3.11)	
Unpaid Family Workers	-.222	.159	-.204	.146	-.242	.167
	(20.4)		(11.0)		(18.1)	
All Working	-.0068	.469	.0519	.591	-.0822	.395
	(.46)		(2.17)		(4.60)	
Men Age 15-64						
Full-Time Employees	.407	.608	.167	.555	.461	.638
	(24.6)		(4.59)		(24.9)	
Part-Time Employees	.0449	.017	.0455	.013	.0453	.019
	(9.94)		(5.22)		(8.34)	
Entrepreneurs	-.388	.245	-.100	.093	-.498	.332
	(27.0)		(4.65)		(27.5)	
Unpaid Family Workers	-.0288	.053	-.0963	.144	-.0037	001
	(3.77)		(3.61)		(2.91)	
All Working	.0351	.922	.0163	.806	.0043	.989
	(3.95)		(.57)		(1.04)	

[a] Wage coefficient from predicted log wage of full-time employees in linear probability function for primary participation in employment type. Elasticity of participation with respect to wages is obtained by dividing wage coefficient by mean participation rate. Also included in the specification is a quadratic in age and an intercept.

Table 7

Labor Force Participation Response to Own Wages of Women and Men by Type of Employment and by Marital Status: 1995

	All Persons		Single		Married	
	Wage Coefficient[a]	Mean Participation Rate	Wage Coefficient[a]	Mean Participation Rate	Wage Coefficient[a]	Mean Participation Rate
Women Age 15-64						
Full-Time Employees	.302	.411	.213	.558	.318	.345
	(26.2)		(11.1)		(23.3)	
Part-Time Employees	.0139	.009	.0331	.014	.0058	.006
	(6.03)		(5.65)		(2.54)	
Entrepreneurs	-.0503	.064	-.0545	.059	-.0502	.066
	(8.23)		(5.09)		(6.75)	
Unpaid Family Workers	-0.15	.100	-.0267	.040	-.129	.126
	(14.1)		(2.94)		(13.1)	
All Working	.161	.583	.162	.672	.144	.543
	(13.4)		(8.74)		(9.80)	
Men Age 15-64						
Full-Time Employees	.278	.633	.266	.696	.281	.609
	(19.9)		(7.00)		(18.5)	
Part-Time Employees	.0366	.021	.0526	.017	.0358	.022
	(8.43)		(4.80)		(7.41)	
Entrepreneurs	-.213	.234	-.0648	.095	-.248	.288
	(17.0)		(2.65)		(16.9)	
Unpaid Family Workers	-.0098	.022	-.0242	.048	-.0068	.013
	(2.16)		(1.32)		(1.83)	
All Working	.0916	.911	.230	.856	.0615	.932
	(11.1)		(7.82)		(7.96)	

[a] Wage coefficient from predicted log wage of full-time employees in linear probability function for primary participation in employment type. Elasticity of participation with respect to wages is obtained by dividing wage coefficient by mean participation rate.

part to their increased education, would have had no effect on total participation (all working) of women in the labor force, because it would have been associated with increased participation among single women and decreased participation among married women. Two decades later, both marital groups are entering into employment outside of the family in response to improvements in their wage opportunities, as observed in most labor supply studies in high-income countries. These regularities in uncompensated wage effects on labor supply are consistent with the evidence presented earlier that household inequality in income per adult is more equal when value is assigned to the time not spent in the labor force. Full income inequality is distinctly more equal than market income inequality in Taiwan by 1995.

8. Summary and Conclusions

Measured change in the distribution of income across households from 1964 to 1995 is sensitive to how income is adjusted for the changing composition of households. Without adjustment, households are all treated as the same in terms of the welfare they can produce with a given income. Inequality then decreased slightly from 1964 to 1975 and increased from 1980 to 1993. This is the general time series pattern reported in the literature on Taiwan's income distribution. But as illustrated in Figures 5 and 6, when household income is divided by household size (or number of adults), inequality across households changes substantially, and time trends toward increasing inequality of household incomes from 1976 to 1995 are eliminated. Arguably economic inequality is better approximated by household income <u>per capita</u> (or per adult). In this case many stylized facts regarding aggregate changes in personal income inequality in Taiwan must be reappraised and inequality trends observed in other developing countries may need to be reexamined.

There is no ideal standardization of household income for household composition because the composition reflects household production and consumption technology, prices, the relative wages of women, men and children, possibly cultural factors, and the heterogenous preferences of household members. In sum, household composition is endogenous, responding to the income earning possibilities open to individuals, and only these latent income earning opportunities summarize the true exogenous state of economic inequality. First, there is the fertility decision which is reflected most clearly by the presence in the household of dependent children, defined here simply as persons under age 15 within a sample of households where the head's age was between 30 and 50. Second, there is a decision to support elderly within the household, who may be dependent on their adult children for consumption of market goods and for the provision of physical care. Third, there is the propensity for prime-aged adults to live together to realize economies of production and consumption, as well as to share public goods, such as children. All three of these processes are complex and examined here in only cursory fashion. But all three would appear to be related to income or expenditure levels, and thus they may possibly not be exogenous across households with regard to the latent distribution of income opportunities. The challenge raised by this paper is how might the social statistician adjust household income for household composition to better approximate the welfare opportunities of its members. Only when this challenge has been met squarely and somehow resolved, can the distribution of economic resources among households be more meaningfully compared across societies or over time within societies.

One simple step is to separate the fertility decision from the household income distribution, by means of analyzing household income per adult. Market income per adult may contribute to couples getting married or not, how many children they have, how much they invest in each of their children that could be closely related to their fertility choice. Household income per adult is thus an improved indicator of income opportunities of adult decision makers in a population and expenditures per adult may proxy more closely the persistent or permanent income concept that would be more relevant to life cycle decisions, such as marriage and fertility. It is shown that both in 1976 and 1995 the estimated relationship between household expenditures (income) per adult and fertility is positive in Taiwan, and that the elasticity has not changed much in these two decades, although the level of fertility has fallen sharply. There is also a clear inverse relationship between household expenditures (income) per adult and the extension of households to accommodate the elderly or other prime-aged adults. Thus, as incomes increase, more of the young split off and the old retain their separate households. Since these extreme age groups tend to have below average incomes, the process of development adds to measured income inequality across households, even after household income is divided by adults to represent more adequately the welfare of decision makers.

However, the economic resources available to adults is not necessarily identical to the household's market income per adult. Time of adults not spent in market labor force activity may increase the economic welfare of household members. This time is not only consumed as leisure, but also may be employed in home production of goods and services, including children. In this time period in Taiwan, fertility is declining and the labor force participation of women is increasing, particularly outside of agriculture and outside of the home or self employment. This reallocation of women's time from the home to the market wage work might mask underlying shifts in economic inequality.

To explore this possibility, an empirical approximation for full income is proposed. I have attempted to impute to all adults 15 to 65 the market earnings they could earn if they were typical full-time wage earners. An exception was made for entrepreneurs if they earned more than this imputed full-time earnings, in which case the entrepreneur's actual income receipts are retained by the household. Also a special allocation rule was adopted for the elderly, over age 64, for whom actual earnings are attributed to the household rather than the imputed full-time earnings. The rationale is that exogenous variation in health status may widely affect the allocation of time by the elderly to market work and therefore those not working may have health limitations. Following this strategy for allocating earned income to households, actual nonearned income and transfers are then attributed to the full-time earnings of households.

Inequality in "full income" per capita is decreasing in Taiwan across households from 1976 to 1995, whether measured by the Gini or variance of log income and whether the imputation of earnings is based on individuals or also couples. The Gini in full income per adult has also decreased marginally in this time period, based on the individual imputed earnings. Inequality in "full income" tends to be smaller than inequality in market income, especially by the end of the period. This suggests that persons with higher wage opportunities are increasingly likely to be working more of their time in the labor market. Thus, inequality in market income is more unequal than inequality in full economic opportunities across households. Stated differently, the poor have more nonmarket time than the rich, a pattern observed to be emerging as well in the United States in recent years (Juhn and Murphy, 1997; Welch, 1997). This pattern is also evident if household income is standardized by all persons, including children, or not standardized at all. The earlier noted positive correlation between market per capita income and fertility is not greatly affected when actual income is replaced by the imputed values of full household income per adult.

Another component of family composition is the extension of the family to share housing with additional generations, or specifically, to have mature children coreside with their elderly parents. The frequency of such extended family arrangements is diminishing in Taiwan as in many other advanced economies. Simple regression analysis indicates that the elderly are less likely to live with their children, if the adult children have more income per adult. Thus, if these cross sectional patterns indicate how coresidential behavior changes over time, the large increase in income levels per adult in Taiwan (Figure 1) should have contributed to the increasing propensity of the elderly to live in separate households. The increased life expectancy of the elderly in Taiwan probably also signals that the elderly are healthier than they were in the past, and thus less limited by health in their capacity to attend to their daily care and consumption needs.

The demographic transition in Taiwan has changed the age composition of the population, increasing the share of the population in between the ages of 30 and 44 and over 60. The increase in income has contributed to the increased likelihood that the elderly are heads of their own households, rather than living with their children. Both the demographic aging and the economic growth of Taiwan has increased the share of households headed by elderly. Many studies suggest that income inequality tends to increase at older ages in a cross sectional survey, and when repeated cross sections are compared it is common to find that income and expenditure inequality also increases statistically within cohorts as they become older (Deaton and Paxson, 1994). But in Taiwan, this is evident only among cohorts between the ages of 35 and 54, and the margin of increase in inequality with aging is modest (Table 3). Nonetheless, the increasingly older age distribution of heads of households has contributed in Taiwan to an increase of a few percent in the variance of household income per adult across all ages. This effect of the changing age composition does not imply that later born cohorts, as they live out their lives, will encounter a substantially greater level of inequality, but only that economic resources of the old are somewhat more unequally distributed than those of the young, and the old are becoming an increasing share of the heads of households in Taiwan. This secular change in the age composition of household heads in Taiwan accounts for about a third to a half of the small increase in aggregate variance of the log of household market income per adult from 1976 to 1995. This effect of the demographic transition is relatively small compared to that contributed by changes in size and composition of households noted earlier, but as the aging of the population continues and the elderly increasingly reside on their own, this effect of the changing age composition of household heads may become a more significant source of change in measured income inequality in Taiwan.

Notes

1. Only a few countries outside of Europe, North America, Oceania and Japan started their demographic transition earlier than Taiwan. Exceptions are Argentina and Uruguay, with their large European settlements, that experienced a decreasing level of fertility before the rest of Latin America.

2. Regressions have been estimated by Chu (1996: Table 3) to explain mean wage payments across (84 and 106) cross tabulated groups defined by education, age, city/noncity residence, and sex that imply wage payment differentials by education and urbanization decreased from 1966 to 1977. The relative wage effect of an additional year of schooling decreased from 16 to 7 percent in this period of only 11 years, suggesting a sharp decline in private wage returns to schooling. By 1976 when the micro data are available, the relative wage returns to schooling for individual full-time male workers are 7 percent and for females 9 percent (See Appendix Table A-3). The relative wage effect of working in a city fell from a premium of 65 percent in 1966 in the aggregate regression of Chu (1996) to one of 31 percent in 1977. Chu concludes from these wage payment regressions across groups that the wage differentials by skill decreased in Taiwan in this period, due to both the increased supply of better educated workers and the increased demand for less-skilled workers in the well-dispersed export industries that rapidly reduced underemployment. Without information reported in the survey on hours worked, it is not clear how the wage variable is "normalized" for labor supply by Chu, and what bias that normalization might introduce into the reported aggregate wage regressions.

3. Kuznets (1962) had earlier noted a similar pattern in the United States where household income inequality increased from 1945 to 1955, and he explained it in terms of an "undoubling" of multigenerational households after World War II when the stock of housing began to catch up with consumer demands that had been rationed by wartime mobilization efforts. In this decade of increasing incomes there was a large decline in the proportion of large households (multigenerational) but little change in household per capita income inequality. The same pattern is noted in Taiwan in the period from 1980 to 1995.

4. In this shorter period of observation, consumption may be preferable to income as a measure of welfare, if households have the capacity to smooth their consumption according to their lifetime budget constraint. The sum of household expenditures on consumption, including imputed values for home produced consumer goods, may thus measure more accurately welfare inequalities than current income receipts.

5. One approximation for average individual welfare divides household income by the square root of the household size, i.e., twice the income is needed to maintain a four-person household at the same welfare level of consumption as the income for a two-person household (Gustafsson, 1995).

6. The closure of the gender gap in schooling is readily shown in these data. In 1976 men and women born between 1917 and 1921 had an average difference in schooling of 4.2 years, whereas in 1995 men and women born between 1966 and 1970 report a gender gap in schooling of only .23 years. The lack of an increase in the wage of Taiwanese women relative to men during this period is interpreted by some as evidence of labor market discrimination against women (Zveglich et al.,1997), but may be also be explained by the increase in female labor force participation that tends to be associated with a decline in the labor market experience of female workers compared with male workers. The rapid entry of relatively inexperienced female workers in the United States during the 1950-1975 period has been similarly offered as an explanation for the stability of female wages relative to male wages in this period, and this hypothesis predicted that female wages would begin to overtake male wages after 1980, when female worker experience was expected to begin to overtake males worker experience (Smith and Ward, 1985). And they did.

7. Instances can be found when transfers from the very rich to still richer households can decrease the variance in log incomes, but this violation of the Pigou-Dalton "transfer principle" is an extreme and hardly a reason to reject the variance of log income as a generally useful summary measure of inequality that is particularly sensitive to inequality in the lower income tail of the distribution.

8. After 1987 one can compare empirical results that are based on all coresident couples. This restriction did not seem important in the patterns discussed in this paper. It is also possible to impute matches of spouses according to age, in those infrequent cases where a married male nonhead of household might be matched with alternative "prospective wives" in the household. In 1990 there were 2,997 household heads and their spouses who were both full-time wage earners, and another 234 full-time wage earning couples who were not heads. In 1995 there were 3,056 head couples and 138 non-head full-time wage earner couples. These groups are later used to infer how "full income" varies across households with couples.

9. From 1976 to 1993 the proportion of elderly men and women (over the age of 64) living with their children (or children-in-law) decreased within various age groups in Taiwan, while the proportion living on their own and remaining heads of households increased, as tabulated in Appendix Table A-2. For example, the proportion of men age 65 to 74 living in a household with their children was 30 percent in 1976 and that proportion had declined more than one-fifth to 23 percent by 1993, while the proportion living on their own as head of household increased from 35 percent to 52 percent. This change in living arrangement could potentially be attributed to improved health among the elderly that allowed more of them to take care of themselves (Schoenbaum, 1995). The economic growth in incomes among the elderly and their children also allowed them to express their preference for the more costly practice of retaining separate living units. Relative prices may also have had a hand in this change.

10. It is later shown that in Taiwan children are not concentrated in lower income households as in the United States, for example.

11. The difficulty with this illustrative specification is that the income variable is divided by adults which imparts a definitional negative correlation between the normalized income variable and the number of prime aged adults and the share of elderly in the household.

12. This is merely the standard decomposition of a variance into between group and within group components, as developed by R. A. Fisher (1930), which separates the variance ascribable to one group of causes from the variance ascribable to other groups. In a multiple factor classification scheme, the share of the variance in income explained by any one factor will not be independent of what other factors are simultaneously included, unless all the conditioning factors happen to be orthogonal (Schultz, 1965).

13. "Other employee income" reported in the survey (not full-time employee or part-time employee compensation) is allocated to full-time and part-time earnings in proportion to their reported direct compensation. This treatment of other employee income is based on the assumption that these income flows are related to bonuses and fringe benefits and although they may be a larger fraction of full-time compensation than part-time compensation, a proportional distribution seemed reasonable.

14. Because in the early years of the survey the household members are not unambiguously linked to their spouse except when they are the head of the household or the head's spouse, these earnings equations for couples who are both full-time earners are also restricted to the heads. This preserves the same criteria for selection from 1976 to 1995. Adding in 1995 the sample of couples who were both full time wage earners but are not heads made no substantial differences in estimated earnings equations or covariances between spouses, or resulting full income estimates in Table 4 and 5.

15. These correlations of spouse education are higher than noted in many other populations. See for example for the United States (Mare, 1981, 1991) and other countries (Shavit and Blossfeld, 1993).

16. Because the dependent variable is binary, these OLS estimates are not as useful for hypothesis testing as logit or probit estimates, and caution should be used in interpreting the reported t ratios in parentheses beneath the regression coefficients. The standard errors from OLS in this case are biased. On the other hand, retaining the linear specification provides a simpler interpretation of the magnitudes of the coefficients on the log wage variable across the various types of employment.

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Table A-1

Coefficient on Household Income or Expenditure per Adult in Regression on Household Composition: 1976 and 1995

	log Income per Adult		log Expenditure per Adult	
	1976	1995	1976	1995
1. Mean Log Income or Expenditure per Adult (Standard Deviation)	3.688 (.501)	5.852 (.538)	3.534 (.480)	5.632 (.539)
2. Log Household Size	-.189 (20.2)	-.202 (25.6)	-.181 (18.3)	-.174 (21.8)
3. Log Number Prime Adults	-.332 (34.7)	-.299 (40.0)	-.373 (37.6)	-.294 (39.0)
4. Log Nonelderly/ Prime Adults	.175 (20.8)	.155 (26.5)	.228 (26.3)	.191 (33.3)
5. Log Households re/Nonelderly	-.0319 (16.3)	-.0578 (18.5)	-.0354 (17.2)	-.0710 (22.9)
6. Number Prime Adults	-1.041 (32.5)	-.875 (42.6)	-1.185 (35.8)	-.878 (42.7)
7. Children/Prime Adults	.0635 (15.5)	.102 (18.0)	.0682 (15.9)	.122 (21.5)
8. Elderly/Prime Adults	-.0635 (15.5)	-.102 (18.0)	-.0682 (15.9)	-.122 (21.5)

Note: Sample includes households whose head is age 30 to 50. Regression of 8 measures of household composition on log income or expenditures per adult in household, controlling for age of head and age squared. Children Age < 15; Prime Adults, 14 < Age < 66; Elderly Age > 65.

Absolute value of t-ratio reported in parentheses beneath regression coefficient.

Table A-2

Distribution of Elderly by Age, Sex and Living Arrangements

	50-64		65-74		75+	
1976	Male	Female	Male	Female	Male	Female
Heads or Spouses of Heads	2051	1069	171	97	36	14
Both Living Together	1774	998	142	79	29	7
Only One Living	277	71	29	18	7	7
Living in Household When Child is Head or Spouse of Head	403	706	263	434	92	194
Both Living Together	332	366	179	124	49	16
Only One Living	71	394	84	310	43	178
Living in a Household Where Child is not the Head or Spouse of Head	27	43	14	43	16	52
Uncertain of Who Make up a Pair						
Total	2935	3647	882	1105	272	468
1993						
Heads or Spouses of Heads	3176	2399	1095	550	272	141
Both Living Together	2893	2171	843	419	166	63
Only One Living	283	228	252	131	106	78
Living in Household When Child is Head or Spouse of Head	1130	1759	941	962	352	461
Both Living Together	1019	1242	770	480	212	78
Only One Living	111	517	171	481	139	383
Living in a Household Where Child is not the Head or Spouse of Head	45	67	56	73	76	156
Uncertain of Who Make up a Pair						
Total	8659	8383	4128	3096	1323	1360

Table A-3

Earnings Function Estimates for Full-Time Employees and Couples Who Both Are Full-Time Employees

Year Explanatory Variable	All Full-Time Employees		Full-Time Employees who are Couples and Heads of Household	
	Male (1)	Female (2)	Male (3)	Female (4)
1. 1976				
Education Years	.0670	.0889	.0626	.0930
	(46.6)	(38.4)	(21.1)	(23.6)
Post Schooling Experience	.0570	.0249	.0154	.0066
	(40.5)	(11.6)	(3.16)	(1.12)
Experience^2(x10^{-2})	-.0911	-.0344	-.0351	.0055
	(30.6)	(6.52)	(3.58)	(.44)
Constant	9.73	9.44	10.33	9.55
	(459.)	(325.)	(141.)	(116.)
R^2	.323	.328	.390	.419
(n)	(7780)	(3354)	(1108)	(1108)
2. 1995				
Education Years	.0775	.102	.0738	.109
	(65.1)	(58.7)	(38.3)	(43.6)
Post Schooling Experience	.0560	.0304	.0402	.0207
	(58.2)	(26.3)	(16.3)	(7.36)
Experience^2(x10^{-2})	-.0878	-.0309	-.0681	-.0111
	(44.7)	(11.4)	(14.7)	(1.87)
Constant	11.60	11.17	11.90	11.19
	(633.)	(445.)	(277.)	(235.)
R^2	.424	.389	.409	.424
(n)	(10354)	(6914)	(3052)	(3052)