Intertemporal Choice and Inequality in Low-Income Countries: Evidence from Thailand, Pakistan, and India*

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Abstract

It is well-known that within-cohort consumption inequality increases with age in developed countries and this pattern is consistent with the permanent income hypothesis, under which households smooth consumption through credit markets in the short-run against transient shocks and in the medium-run over the life cycle. This paper provides evidence regarding the age effects in within-cohort inequality for several low-income developing countries, where credit markets are underdeveloped. Empirical results show patterns unnoticed in the literature. The within-cohort inequality in consumption often decreases with age, and the divergence from the pattern from those observed in developed countries is larger among uneducated and rural households. We provide an interpretation that the decreasing age effect in consumption inequality within cohort, found widely in low-income regions and classes in Asia, is consistent with the autarky model with decreasing within-cohort inequality in income in agricullture-based rural societies.

1 Introduction

It is well-known that within-cohort consumption inequality increases with age in developed countries and this pattern is consistent with the permanent income hypothesis, under which households smooth consumption in the short-run against transient shocks and in the medium-run over the life cycle. The seminal paper by Deaton and Paxson (1994) showed that the inequality indeed increased with age at a similar speed in US, Great Britain, and Taiwan. They interpreted this as the reflection of cumulative difference in the effects of luck on consumption. Storesletten et al. (2004) extended their analysis for the US case, showing that age effects in income and consumption inequality within cohort are indeed consistent with theoretical prediction of an overlapping generations general model in which households face uninsurable earnings shocks over the course of their lifetime. An interesting finding by Storesletten et al. (2004) is that the risk faced by households that realized throughout the working years is more important than the risk that realized before entering the labor market.

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In low-income developing countries, people in general, and particularly less-educated, poor farmers, have few means to hedge against the vagaries of income shocks that may put their livelihood at risk (Fafchamps 2003; Dercon 2005). This implies that the risk realized throughout the working years should affect consumption dynamics of households in these countries more substantially. On the other hand, in these low-income countries, the development of credit markets has been lagging behind the development of output and other factor markets (Kochar 1997a; 1997b). This means that poor households in these countries have difficulty in smoothing consumption intertemporally through credit markets, while they may utilize informal measures of risk-sharing among fellow villagers to smooth consumption across states (Townsend 1994). Regarding shocks to income, residents in low-income developing countries are often dependent on farming, whose returns may be more subjective to transient shocks than to permanent productivity shocks as in developed countries (see Storesletten et al. 2004 for the US case).

Under these conditions, we expect the within-cohort income/consumption inequality in low-income developing countries to show an age pattern that is different from those in developed countries. The shape of the age effect in inequality will give us important information regarding the intertemporal choice available for residents in low-income countries. In spite of this, there are very few studies that directly extended the empirical exercises of Deaton and Paxson (1994) to these countries.

Given this paucity in the literature, this paper attempts to accumulate evidence regarding the age effects in within-cohort inequality for several developing countries, in the spirit of Deaton and Paxson (1994). For this exercise, repeated cross-section datasets of household consumption expenditures are required, covering as many years as possible. As datasets that satisfy this requirement, we employ those from Thailand, Pakistan, and India. To increase statistical efficiency as well as to control for changes in household demographics, we extend the cohort-level regression model by Deaton and Paxson (1994) to a model at the household level. Empirical results show patterns unnoticed in the literature. The within-cohort inequality in consumption often decreases with age, and the divergence from the pattern from those observed in developed countries is larger among uneducated and rural households.

The remainder of the paper is organized as follows. The next section explains an empirical model to estimate the age effect in within-cohort inequality. The datasets are described in Section 3, while the empirical results are presented in Section 4. A speculative discussion to associate the findings of this paper with theories of intertemporal choice is given in Section 5.

2 Empirical Specification

Let c_{igt} be the per-capita real consumption expenditure for individual/household i belonging to cohort group g for the survey round t. We define cohort by the birth year of the household head. We treat i as the individual or the household interchangeably in theoretical discussion while we treat i as the household weighted by its number of household members in empirical analyses since all expenditure data are at the household level.

Given cross-section data of c_{igt} for year t, we can calculate various measures of inequality within cohort by aggregating c_{igt} across i belonging to g. Following the literature, we employ the variance of log consumption as the main measure of inequality in this paper, i.e., $Ineq_{gt} = Var_{i \in g}(\ln c_{igt})$.

Deaton and Paxson (1994, Figures 1-3) first plot the observed values of $Ineq_{gt}$ for the same g but in different t on the axis of the age of the household head. These plots show an increasing age effect. They then estimated the cohort-level model:

$$Ineq_{gt} = Var_{i \in g}(\ln c_{igt}) = \sum_a \alpha_a Age_{gt} + \sum_g \beta_g Cohort_g + u_{gt}, \tag{1}$$

which is basically a regression of the cohort-level inequality on age fixed effects (α_a) and cohort fixed effects (β_g). Fitted values of the age fixed effects, $\hat{\alpha}_a$, show the dynamics of within-cohort inequality across age and their confidence intervals can be estimated from standard errors of $\hat{\alpha}_a$. Fig.4 of Deaton and Paxson (1994) shows the age effects thus compiled, which are indeed increasing with age at a similar speed in US, Great Britain, and Taiwan.

With the same spirit as Deaton and Paxson's (1994), we estimate a household-level model:

$$(\ln c_{igt} - \bar{\ln c}_{gt})^2 = \sum_a \alpha_a Age_{igt} + \sum_g \beta_g Cohort_{ig} + X_{igt}\gamma + u_{igt}, \tag{2}$$

where $\bar{\ln c}_{gt}$ is the cohort average of $\ln c_{igt}$ for $i \in g$ in year t, X_{igt} is a vector of demographic variables that characterize household i in year t and region fixed effects. By adopting this specification, we can expect a gain in statistical efficiency as well as we can control for changes in household demographics in a more straightforward way. Similarly to Deaton and Paxson (1994), fitted values of the age fixed effects, $\hat{\alpha}_a$, show the dynamics of within-cohort inequality across age and their confidence intervals can be estimated from standard errors of $\hat{\alpha}_a$.

In estimating either (1) or (2), it should be noted that macroeconomic shocks (which can be represented by survey-round fixed effects) are already controlled in an implicit way through the combination of the cohort fixed effects and the age fixed effects. It is well-known that age, cohort, and year fixed effects are perfectly collinear so that it is not possible to

identify all of them in a linear model. When cohorts in different rounds in the repeated cross-section dataset are well-overlapping (i.e., the repeated cross-section dataset with a long time horizon), age effects are estimated reliably, while the parameter estimates for the age effects become unstable in the repeated cross-section dataset with a short time horizon because the identification is based on a smaller number of overlaps of cohorts.

When household income data are available in a repeated cross-section dataset, we calculate y_{igt}, per-capita real income for individual/household i belonging to cohort group g in year t in the similar manner. We then estimate equation (2) replacing $\ln c_{igt}$ by $\ln y_{igt}$ and replacing $\bar{\ln} c_{gt}$ by $\bar{\ln} y_{gt}$.

3 Data

3.1 Thailand

The data source for Thailand is the *Household Socio-Economic Survey* (Thai SES data hereinafter). Thai SES is conducted by the National Statistical Office of the Government of Thailand. Since 1998, the survey has been conducted every year. A nationally representative sample is drawn each time and surveyed using a detailed questionnaire on household demographics, income, and consumption, covering approximately 11,000 to 35,000 households.

Ten rounds of Thai SES spanning a period of 19 years (1986, 1988, 1990, 1992, 1994, 1996, 1998, 2000, 2002, 2004) are employed in this paper. The period covered by Thai SES corresponds to a period of the "Asian miracle" from the 1980s to the mid 1990s, then the Financial Crisis in 1997, and finally a recovery from it (Kurita and Kurosaki 2007). Thai SES data have a good number of overlapping cohorts across rounds. This is a great advantage in estimating age effects following equation (1) or (2).

Real per-capita consumption (c_{igt}) was calculated by dividing total household consumption expenditure by the number of household members and the government price index. Another advantage of Thai SES is that it includes information on household income. By comparing the dynamics of within-cohort consumption inequality with that of within-cohort income inequality, we can obtain insights regarding factors governing the intertemporal consumption choice of Thai consumers.

3.2 Pakistan

The data source for Pakistan is the *Pakistan Integrated Household Survey* (PIHS) in 1998/99 and 2001/02 and the *Pakistan Social and Living Standards Measurement Survey* (PSLM) in 2004/05 and 2005/06. PIHS/PSLM are conducted by the Federal Bureau of Statistics, the Government of Pakistan. While PIHS was a survey of income and expenditure, PSLM

in general focused on various aspects of well-being, including income, education, health, etc. The subsample of PSLM were asked additional questionnaires corresponding to the household income and expenditure surveys. This part provides information comparable to PIHS.

In each survey, a nationally representative sample was drawn in two stages: primary sampling units (PSU) with different sampling probabilities were randomly chosen in the first stage; twelve (or eleven in the 2005/06 survey) households were randomly chosen from each PSU in the second stage. The sample size for our analysis is approximately 15,000 households in all four surveys, containing approximately 105,000 individuals.

In the PIHS/PSLM dataset, nominal consumption expenditure including the imputed values of in-kind transactions per capita[1] in Pakistan Rupees is calculated and then converted into a real term by dividing by the official poverty line. This is the concept known as the "welfare ratio." Since the log variance is employed as the measure of inequality, the unit of measurement does not matter.

The period covered by PIHS/PSLM experienced both increase and decrease of poverty (Kurosaki 2009). The average consumption declined from 1998/99 to 2001/02, followed by increases in the next two periods. Its level in 2005/06 is only 15% higher than the 1998/99 level but 21% higher than the 2001/02 level. The movement of average consumption is closely related with agricultural production in Pakistan.

3.3 India

For India, we employ data from NSS expenditure surveys, conducted by the National Sample Survey Organization, the Government of India. This paper reports results based on four rounds with so-called "thick" sample: 1983 (38th NSS), 1987/88 (43rd NSS), 1993/94 (50th NSS), and 1999/2000 (55th). The four rounds thus cover about eighteen years. The NSS expenditure dataset contains incredibly detailed information on consumption items.

A great advantage of these NSS datasets is their large sample size. Each round contains approximately 110,000 sample households. This enables us to keep a sufficient number of sample households when we estimate age effects for sub-group levels. On the other hand, NSS surveys do not include information on household income. Therefore, we estimate age effects in within-cohort inequality only for consumption.

The period covered by the four rounds of NSS experienced a continuous decline in poverty. The average per-capita consumption increased continuously during this period,

[1]To be precise, "per capita" means "per adult equivalence unit," which is the unit adopted by the Government of Pakistan to establish the official poverty line. Individuals who are 18 years old or above are assigned the weight of 1.0 and others are assigned that of 0.8.

resulting in a decline of poverty head count index from 45% in 1983 to 26% in 1999/2000.[2] However, the level of the poverty measures in 1999/2000 has been debated intensively because consumption modules, especially the recall period, were changed in 1999/2000, resulting in the non-comparability problem. Adjustment to solve the non-comparability problem (like the one implemented by Tarozzi 2007) is left for further exercise.

4 Empirical Results

4.1 Thailand

Figure 1-1 shows the age effects in the within-cohort variances of logarithms of consumption/income in Thailand. They are estimated using the individual-level regression model of (2) and ten SES rounds from 1986 to 2004. First, the within-cohort consumption inequality increases with age only during very young periods and then gradually declines until retiring ages. This is in sharp contrast to the increasing age-effect reported for developed countries (Deaton and Paxson 1994; Storesletten et al. 2004). The 95% confidence interval, however, shows that the within-cohort inequality level after the age of mid thirties until the age of late fifties are not significantly different from the inequality level during the early twenties.

On the other hand, the within-cohort inequality in income increases with age increases sharply during very young periods and then remains constant at that level until retiring ages. The 95% confidence interval shows that the within-cohort inequality level after the age of late twenties until the age of late fifties are significantly higher than the inequality level during the early twenties. This is somewhat similar to the patterns reported for developed countries (Deaton and Paxson 1994; Storesletten et al. 2004), although the inequality seems to stop rising at the life-cycle stage much earlier in Thailand than in developed countries.

Figure 1-2 compares the age effects in within-cohort consumption inequality across regions: urban versus rural. An interesting contrast is found. In urban areas, the within-cohort inequality in consumption increases with age steadily until the age of sixties, except for the stagnation in the age of early forties, while in rural areas, the age-effect shows a long period of decline after an initial rise, resulting in a hump shape. In urban areas, the inequality level in the age of late fifties is significantly higher than that in the age of mid twenties, while in rural areas, the opposite holds. The deviation from the pattern observed in developed countries is thus clearer in rural areas. The urban-rural contrast in the age effects exists for within-cohort income inequality as well (not reported). If we plot age effects in income and consumption inequality within cohorts for urban areas only, the shape is quite similar to those reported for developed countries: income inequality increases faster than consumption

[2]Official estimates by the Planning Commission, Government of India.

inequality, suggesting consumption smoothing over the life cycle among urban consumers.

One problem in interpreting the urban-rural contrast is potential endogeneity of migration decisions. On the other hand, the educational level of household heads is pre-determined for households when they decide on their (dynamic) consumption choices. Figure 1-3, therefore, compares the age effects in within-cohort consumption inequality across education levels. Sample households are divided into a high education group (household heads' years of schooling were more than or equal to 7 years) and a low education group (household heads' years of schooling were less than or equal to 6 years). The contrast is sharper than that in Figure 1-2. The shape for the high education group is similar to that for the urban households while the shape for the low education group shows a clearly declining trend without a hump. However, since the standard error is also large for less educated households, the within-cohort inequality level in consumption is not very significantly smaller than that at the age of 19; it is significantly smaller only after the age of late fifties.

The education contrast in the age effects partially reflects the shape of the age effects in within-cohort income inequality (Figure 1-4). To make the consumption easier, the horizontal axis of Figure 1-4 is set equal to that of Figure 1-3. Among the more educated households, the life-cycle shape of income inequality shows a steady increase, whose age slope is steeper than that of consumption inequality. This is similar to observations from developed countries (Deaton and Paxson 1994; Storesletten et al. 2004). On the other hand, the life-cycle shape of income inequality shows a slight decrease among the less educated households, although the decline is not statistically significant, since the confidence interval is wide. The wide confidence interval seems to suggest that household income among less educated households is subject to larger variation across households that is not explained by household size and fixed effects of age and cohorts.

The regional and class contrasts shown in these figures are robustly found. First, using individual data and empirical specification (2), we estimated a model without the household size variable,[3] or different specifications of age and cohort fixed effects in terms of intervals, or a specification using a different threshold to divide households into the more and the less educated. We found little change in shapes of the age effects.

Second, we applied the cohort-level regression model (1) to the ten SES rounds from 1986 to 2004 using various inequality measures. In Figure 1-5, we plot results for three inequality measures: Gini coefficient, General entropy measure with parameter 0, and Atkinson's (1970) inequality measure with parameter 1. As shown in this figure, the education contrast becomes sharper when cohort-level data are used: the within-cohort consumption inequality increases

[3]Robustness check with respect to the expansion of variables X_{igt} is left for further exercise. See the results for Pakistan regarding the robustness check in this direction.

with age among the more educated households while it decreases among the less educated households, and the difference is statistically significant. The choice of inequality measures or the choice of empirical models does not affect the results: within-cohort inequality in consumption is decreasing with age among uneducated households.

Regarding the choice of inequality measures, the sensitivity of the analysis with respect to the inequality aversion parameter might be of interest. Figure 1-6 thus shows the case using different parameters for Atkinson's (1970) inequality measures. As shown by Ligon (2008), there is a one-to-one mapping between Atkinson's inequality measures and the constant relative risk aversion utility function, where Atkinson's inequality aversion parameter increases as the coefficient of relative risk aversion increases. As shown in panel (A) of Figure 1-6, the education contrast in age effects in consumption inequality becomes sharper when a higher value of inequality aversion parameter is used. However, some of the change may simply reflect the difference in units rather than the difference in shape. Therefore, in panels (B) and (C), the vertical axis is adjusted using an index (the predicted value of the reference age is set at 100). Since the associated confidence interval is also larger for cases when a higher value of inequality aversion parameter is used, the overall significance levels of the education contrast is similar regardless of the choice of inequality aversion parameters. The difference due to the choice of inequality aversion parameters lies in the age around fifties and sixties among the less educated households: inequality continued to decline if a higher value of inequality aversion is assumed (panel (C)) while it remains somewhat constant if a lower value of inequality aversion is assumed (panel (B)). This seems to suggest that the decrease in consumption inequality among the less educated households occurs more in extreme values of consumption rather than in moderate values of consumption when households are aged.

4.2 Pakistan

To investigate the shape of age effects in the within-cohort consumption inequality in Pakistan, equation (2) is estimated using four rounds of PIHS/PSLM from 1998/99 to 2005/06. The Pakistani microdata are associated with uneven intervals and are subject to heaping due to measurement errors in the age and the year of birth reported by the household head. To solve these problems, we employ two-year intervals in defining age and cohort fixed effects. The age fixed effects are based on reported ages (odd ages are rounded-up to the next even age) while the cohort fixed effects are defined on two-year intervals associated with the nearest even intervals in survey years.[4]

[4]Namely, we associated year 1999 with the 1998/99 PIHS, year 2001 with the 2001/02 PIHS, year 2003 with the 2004/05 PSLM, and year 2005 with the 2005/06 PSLM and calculated cohort fixed effects based on these years, with two year intervals. The robustness of our results with respect to this specification is discussed below.

Figure 2-1 shows the age effects in the within-cohort consumption inequality, first for all Pakistan and then distinguishing urban and rural areas. First, the point estimates show an urban/rural contrast similar to the one found in Thailand: the urban pattern is a slight increase while the rural pattern is a flat or slightly decreasing one.[5] This seems to suggest that in Pakistan, urban households' intertemporal choice is quite similar to those in developed countries while households in backward rural regions behave in a different way. Second, nevertheless, the age effects are statistically insignificant. In urban areas, the confidence intervals are so wide that the null hypothesis of no change in the within-cohort inequality is not rejected. The wide confidence intervals among urban households suggest that their consumption is subject to a large variation that cannot be explained by age/cohort fixed effects and the household size. In rural areas, the decline in within-cohort inequality is small in absolute values so that they are statistically insignificant.

We then divide sample households by the education status of the household head: a high education group (households with heads who ever entered formal schooling) and a low education group (households with heads who never entered formal schooling), whose results are plotted in Figure 2-2. As in Thailand, the positive slope is observed only among educated households, which is marginally statistically significant. Among the uneducated households, point estimates of age effects show a declining trend, which is statistically significant as well, although marginally. The education contrast in Figure 2-2 shows not only steeper slopes in age effects but also more balanced sizes of confidence intervals than the regional contrast in Figure 2-1.

The education contrast was found robustly for Pakistan as well. In Figure 2-3, the sensitivity of our results with respect to household-level control variables X_{igt} in equation (2) is investigated. Figure 2-2 using the household size as the only household-level control is re-produced in panel (A) of the figure. Although the household size is statistically significant in most of the regression results, its elimination does not affect much the shape of the figure, as shown in panel (B) of Figure 2-3. When we added more controls, namely, linear terms of five demographic variables and the regional fixed effects (province dummies and urban/rural dummies), the education contrast becomes less significant, although the point estimates still show that the age effect in within-cohort consumption inequality is positive among educated households while it is negative among uneducated households (panel (C) of Figure 2-3). Unfortunately, in panel (C), two confidence intervals are overlapping, making it difficult to reject the null that the age effects are the same between educated and uneducated

[5]This contrast is confirmed when we compare the urban regions in Sindh and Punjab (two provinces out of four in Pakistan, known as more developed regions) versus the rural households in Sindh. The pattern in rural Sindh is a slight decrease while the one in urban Punjab and Sindh is a steady increase. See also Kurosaki (2009) for similar contrasts between urban Sindh/Punjab and rural Sindh.

households. Since many of the variables added in X_{igt} for panel (C) of Figure 2-3 are presumably endogenous to households' decision making, we prefer specification (A) or (B) to (C).

The finding that the age effect in within-cohort consumption inequality is positive among educated households while it is negative among uneducated households was confirmed when age effects were identified against round fixed effects instead of against cohort fixed effects (not reported). Different intervals to define age and cohort fixed effects were also attempted to examine the robustness of our results. The results using one year intervals for both age and cohort fixed effects (corresponding to the actual survey years), or three year intervals for both age and cohort fixed effects (corresponding to the nearest even intervals in survey years) resulted in qualitatively the same results. We also attempted other thresholds to classify households into more and less educated households. Under the threshold used to generate Figures 2-2 and 2-3, sample households are divided into two halves with very similar size. Among those household heads who ever entered formal schooling, about one fifth dropped out from primary school without completing the first stage of primary education. When we use the threshold line of "primary (fifth grade) or better" versus "less than fifth grade completion," the results are very similar to Figures 2-2 and 2-3, with similar statistical inference. When we further upgrade the threshold line at "more than primary (fifth grade)" versus "equal to or less than fifth grade," the education contrast become insignificant.

Using the PIHS/PSLM data, age effects in the variances of logarithms of income were also estimated. To our disappointment, the estimated age effects are very unstable, yielding different shapes depending on specifications. Not only the signs of trends changed depending on specifications, but the absolute size of the age effects also changed by a factor of ten. It seems that income data in PIHS/PSLM suffer from frequent misreporting or nonreporting. Because of this reason, age effects of within-cohort income inequality in Pakistan are not reported in this paper.

The education contrast remains intact when we estimate the cohort-level regression model (1) using the cohort-level data calculated from the four PIHS/PSLM rounds. However, the contrast holds for the point estimates only. The estimated confidence intervals are wide for both educated and uneducated households except for a few inequality measures for which the increasing inequality among the educated is marginally significant at a later stage of life. Figure 2-4 reports the results when various parameters are used for Atkinson's inequality measures. Age effects among educated households show a slight increase while those among uneducated households show no trend. In all three cases plotted in Figure 2-4, confidence intervals are so wide that no statistical inference is possible.

4.3 India

Figure 3-1 shows the age effects in the within-cohort variances of logarithms of consumption in India, focusing on the urban/rural contrast, estimated by equation (2). Similar to the case of Pakistan, the Indian microdata are taken from four rounds with uneven intervals and are subject to heaping due to measurement errors in the age/birth year reported by the household head. When we attempted age and cohort fixed effects defined on one year intervals, the resulting age effects were severely affected by the heaping. When we employed two-year or longer intervals, the effects of heaping were mitigated. Therefore, we report results estimated with age fixed effects defined on two-year intervals and cohort fixed effects defined on five-year intervals associated with the nearest even intervals in survey years.[6] When we use all households, the within-cohort inequality is slightly decreasing and the inequality level in the later ages is significantly smaller than that in the age of early twenties. This is similar to the finding in Thailand and again in sharp contrast to the increasing age-effect reported for developed countries (Deaton and Paxson 1994; Storesletten et al. 2004). Across regions, the declining age effects on within-cohort inequality are strongly observed among rural households. Among urban households, the within-cohort inequality declines with age but at a slower pace, while among rural households, it decreases rapidly with a relatively narrow confidence intervals. The deviation from the pattern observed in developed countries is thus clearer in rural areas, as in Thailand. The confidence intervals are narrower among rural households than among urban households, which is similar to the Pakistani case but opposite to the Thai case.

As a more exogenous variable to classify sample households, Figure 3-2 shows the age effects in the within-cohort consumption inequality, focusing on the education contrast. Sample households are divided into a high education group (household heads' years of formal schooling were more than one: "ever" been to school) and a low education group (household heads had never been enrolled in formal schooling). This threshold is the same as the one adopted for the Pakistani case. As in the cases of Thailand and Pakistan, the education contrast is sharper than the regional contrast in Figure 3-1. The confidence interval is also narrower when we use the education contrast. Because of this, unlike the case of Thailand, the decreasing age effects among less educated households are highly significant. Thanks to the very large sample size of India's NSS, we are thus able to show a case with statistically significant decline in age affects in within-cohort inequality.

The regional and class contrasts shown in these figures are robustly found. First, using

[6]Namely, we associated year 1984 with the 38th NSS, year 1989 with the 43rd NSS, year 1994 with the 50th NSS, and year 1999 with the 55th NSS and calculated cohort fixed effects based on these years, with five year intervals.

individual data and empirical specification (2), we estimate models with different specifications of age and cohort fixed effects in terms of intervals or a specification using a different threshold to divide households into the more and the less educated. Regarding the former, the use of one year intervals for both age and cohort fixed effects resulted in a severe heaping problem, as already noted. On the other hand, other intervals worked well: we attempted age fixed effects in two or three or five years intervals; cohort fixed affects with five-year intervals associated with the nearest even intervals in survey years, or, fixed effects with three-year intervals using the actual years of survey. The results were very similar to those reported in Figures 3-1 and 3-2. Regarding the latter, the education contrast becomes less sharp (but still statistically significant) when households educated more than primary (fifth grade) are compared with others. This observation is qualitatively similar to that found in Pakistan. Since a substantial number of household heads had never entered formal schooling in India, the threshold reported in Figure 3-2 is a more natural choice.

Second, we estimate the cohort-level regression model (1) using the cohort-level data calculated from the four NSS rounds. This approach has an advantage of using various inequality measures. The results robustly support the shape of Figures 3-1 and 3.2, regardless of the choice of inequality measures.[7] As an example, panel (A) of Figure 3-3 shows the cohort-level regression results when Atkinson's inequality measure with parameter one is employed. The within-cohort inequality is associated with little age effects among the educated households while it is associated with significantly negative age effects among the uneducated households. As in Thailand, the decline of inequality among the less educated becomes sharper and steeper when the inequality aversion parameter is increased (compare panel (B) and (C) of Figure 3-3). The confidence intervals are reasonably narrow even when we use the cohort-level data. This is in sharp contrast to the case of Pakistan. Both Indian and Pakistani datasets come from four rounds of repeated-cross section microdata. However, the period covered by the Indian data is much longer (eighteen years) than that by the Pakistani data (seven years). Because of this difference, the overlapping of cohorts is thicker in the Indian data, resulting in better identification of age effects.

5 Theoretical Interpretations

The results in the previous section show that within-cohort consumption inequality does not increase as households get older in various cases from developing countries. The deviation from the increasing pattern is especially significant among less educated households. This finding thus suggests that the applicability of the permanent income hypothesis with perfect

[7]Details are available on request. See also the results reported by Kurita (2005), who found a very strong contrast depending on the education levels of household heads using Gini coefficients.

credit markets (PIH-UC) is not very high for these cases. Theoretically, several alternative models of intertemporal choice are proposed in the literature.

First, as a benchmark, the model of full risk-sharing (or Arrow-Debreu complete markets model) achieves more efficient resource allocation than the PIH-UC model. Under this model, idiosyncratic shocks to income are fully insured so that they should not affect consumption at all (Townsend 1994). This implies that the consumption inequality remains flat, regardless of the dynamics of income inequality.

Another useful benchmark is the model of autarky with no saving technology. Under this model, households always consume their income without any intertemporal or interstate resource allocation. Therefore, the consumption inequality dynamics should closely track the income inequality dynamics.

Between the two polar cases of benchmarks lies the model of PIH-UC. Under this model, idiosyncratic shocks to income affect consumption partially so that the consumption inequality is moderately increasing, when income inequality dynamics shows an increasing pattern (Deaton and Paxson 1994). By incorporating credit constraints (e.g., Deaton 1991), the PIH-UC model can be extended in the direction of the autarky model. Between the two polar cases of the full risk-sharing model and the autarky model also lie models of risk-sharing under private information (Ligon 1998) and risk-sharing with limited commitment (Ligon et al. 2002). Another dimension that should affect consumption dynamics is the possibility of mis-specification of preferences. As Deaton and Paxson (1994) pointed out, the prediction that the inequality in consumption increases with age is based on the assumption that preferences across individuals and across the family cycle are homogenous. When heterogeneity is allowed, we cannot obtain an unambiguous prediction regarding the relation between inequality and age.

The decreasing age effect in consumption inequality within cohort, found widely in low-income regions and classes in Asia, seem to be consistent with the autarky model when income inequality within cohort is decreasing with age. These households are poorer on average than educated or urban households, with limited opportunity to spread risk over time or space. This suggests that the intertemporal choice by these households are better characterized by autarky rather than the permanent income hypothesis or the full risk-sharing hypothesis. The decreasing income inequality could occur in agricullture-based rural societies — since agricultural income is subject to weather and other exogenous shocks, most of which are transitory, the accumulation of permanent shocks in productivity is negligible; on the other hand, when a household is young in its life cycle, agricultural skills are more uncertain (associated with larger variance of transient farm income) and the household head may try several employment opportunities outside agriculture. These may lead to larger

heterogeneity in income earnings when the household head is young.

However, this is only a conjecture. We first need to collect more evidence for the decreasing within-cohort consumption inequality. The evidence shown in this paper is not strong enough, suffering from statistical insignificance in several cases when the number of survey rounds is small or the length of period covered by the surveys is short. After the evidence is accumulated, more rigorous tests to identify the relevant model will be conducted. In such tests, the appropriateness of alternative models, such as risk-sharing under private information, risk-sharing under limited commitment, and more flexible specifications of preference, should also be considered. Rigorous modeling and statistical tests for the appropriate model of household's consumption dynamics are left for further research.

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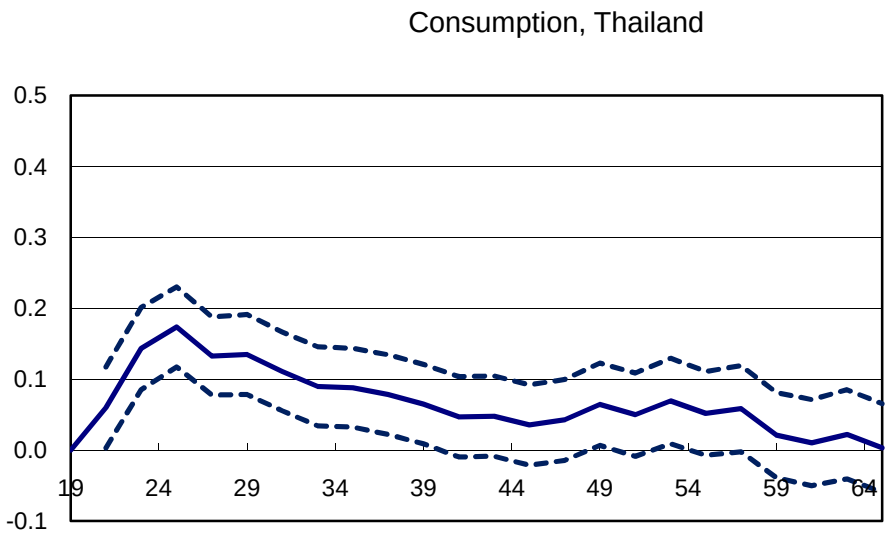
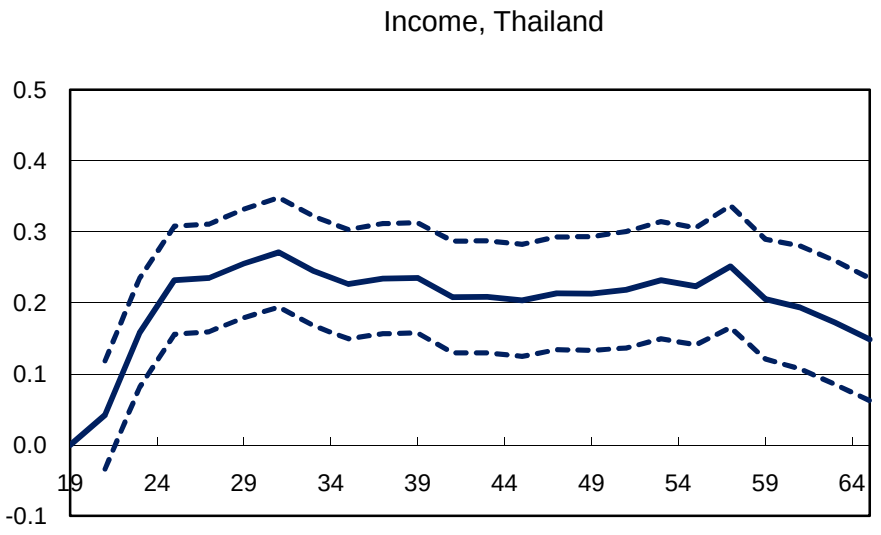
Figure 1-1. Age effects in the variances of logarithms of consumption/income in Thailand



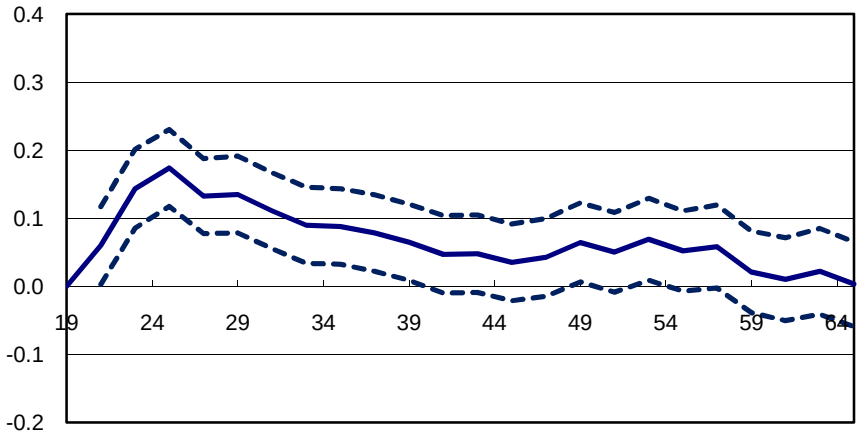


Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Coeff. on AGE f.e. (Age=19 as reference)
Source: Estimated from Thai SES data using weighted regression to control for the difference in sampling probability.
Individual-level data, hh head's age in the range from 19 to 69.
Use 10 rounds from 1986 to 2004, age and cohorts in 2-year intervals
Pooled regression with COHORT f.e. and AGE f.e.
Other controls: demographic variable (hhsize).
NOB: 192,067.

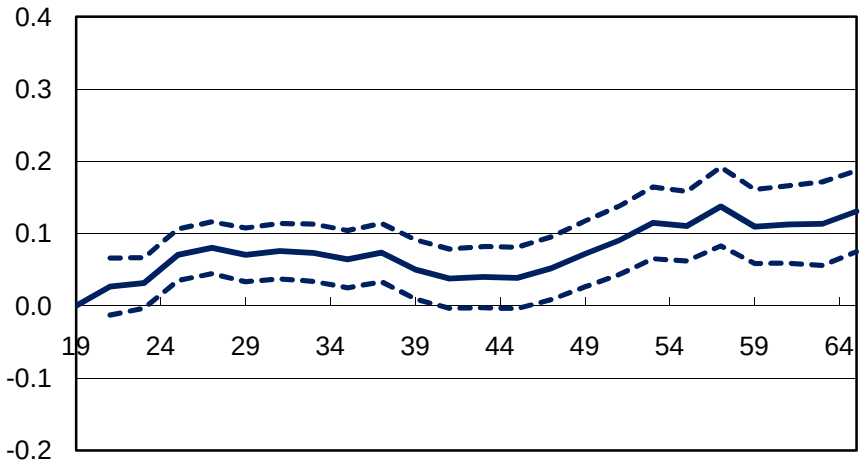
Figure 1-2.
Age effects in consumption inequality and regions, Thailand

Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Coeff. on AGE f.e. (Age=19 as reference)
Source: Estimated from Thai SES data using weighted regression to control for the difference in sampling probability.
Individual-level data, hh head's age in the range from 19 to 69.
Use 10 rounds from 1986 to 2004, age and cohorts in 2-year intervals.
Pooled regression with COHORT f.e. and AGE f.e.
Other controls: demographic variable (hhsize).
NOB: 113,735 for urban areas; 78,332 for rural areas.

All Thailand



Urban Areas



Rural Areas

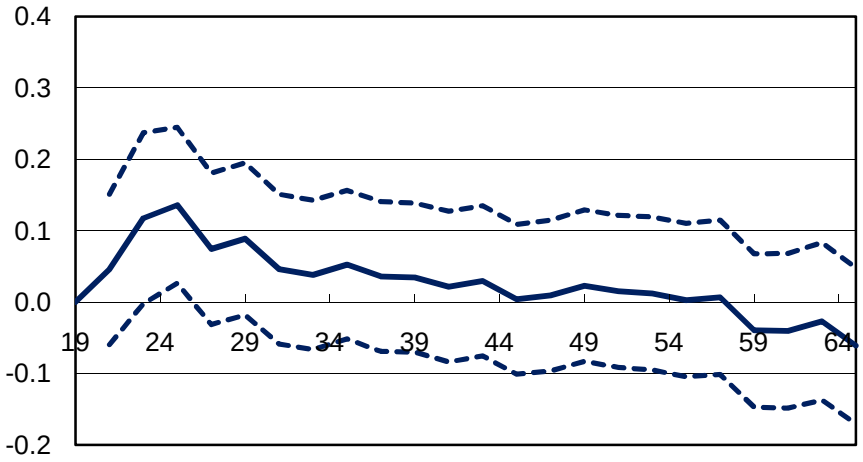
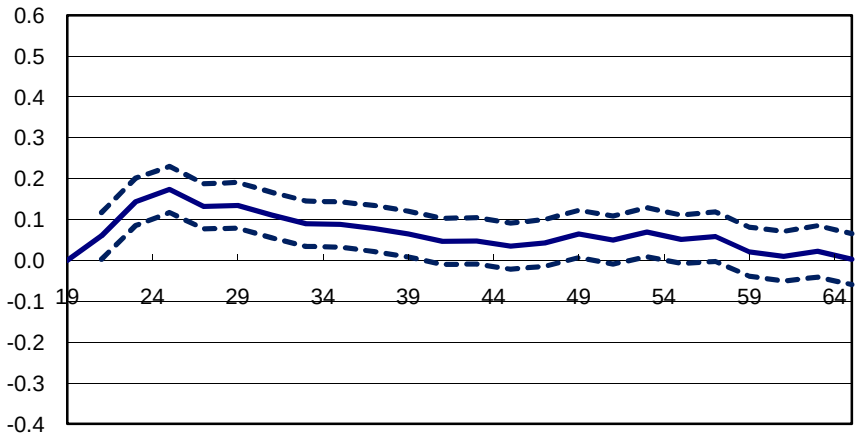
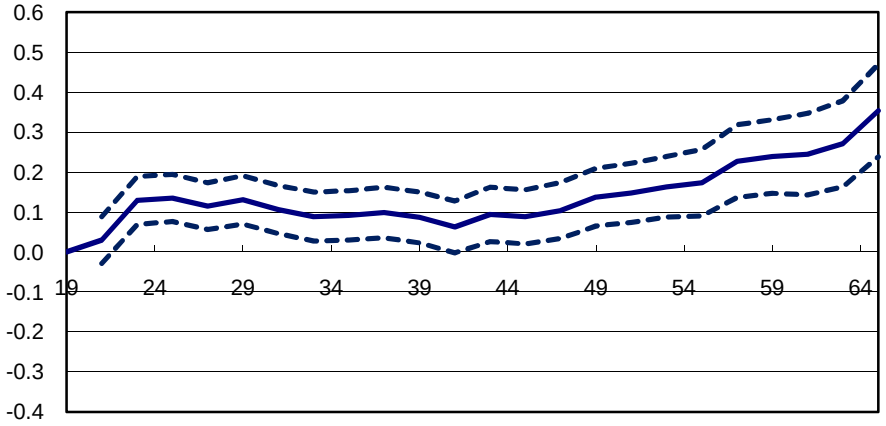


Figure 1-3.
Age effects in consumption inequality and education, Thailand

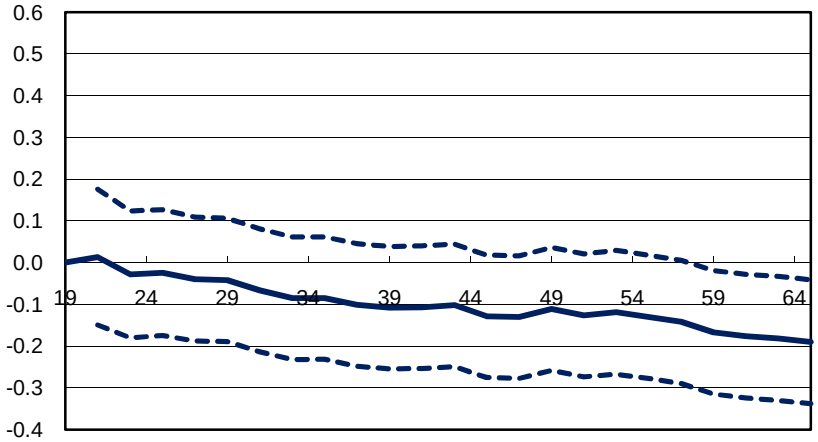
All Thailand



Household head with higher education

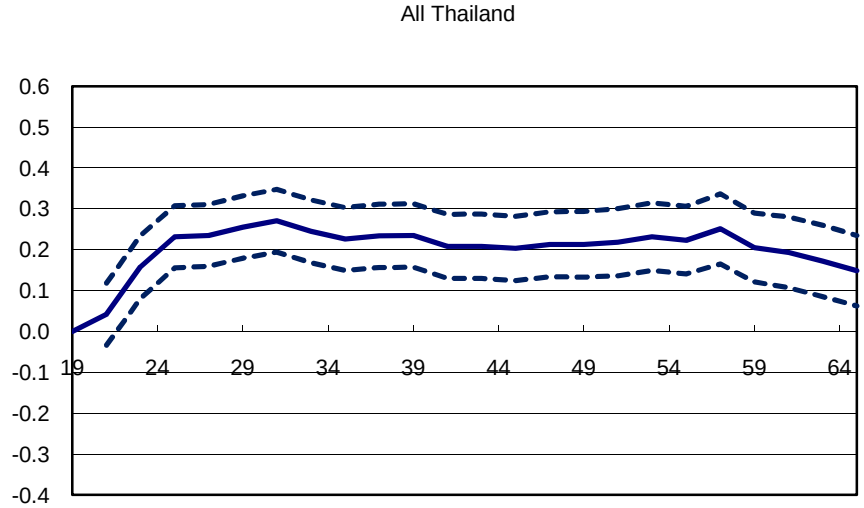


Household head with lower education

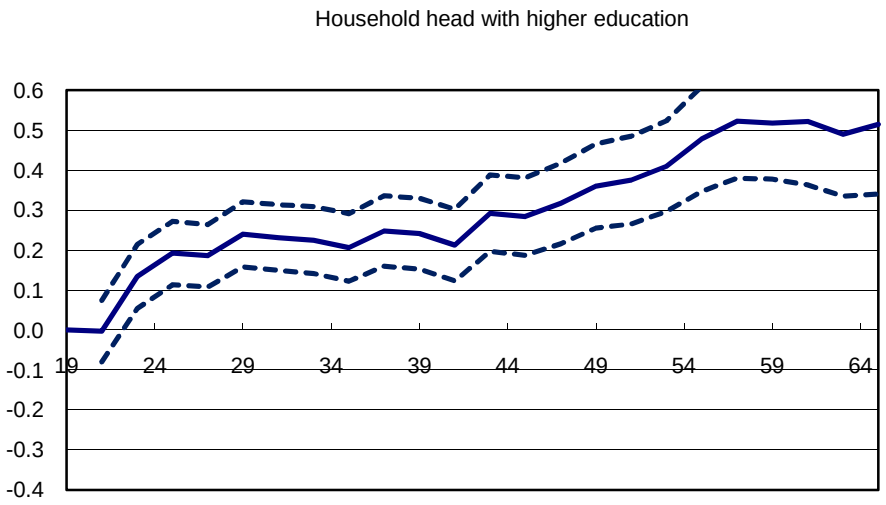


Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Coeff. on AGE f.e. (Age=19 as reference)
Source: Estimated from Thai SES data using weighted regression to control for the difference in sampling probability.
Individual-level data, hh head's age in the range from 19 to 69.
Use 10 rounds from 1986 to 2004, age and cohorts in 2-year intervals.
Pooled regression with COHORT f.e. and AGE f.e.
Other controls: demographic variable (hhsize).
NOB: 76,900 for the more educated; 114,943 for the less educated.

Figure 1-4.
Age effects in income inequality and education, Thailand



Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Coeff. on AGE f.e. (Age=19 as reference)
Source: Estimated from Thai SES data using weighted regression to control for the difference in sampling probability.
Individual-level data, hh head's age in the range from 19 to 69.
Use 10 rounds from 1986 to 2004, age and cohorts in 2-year intervals.
Pooled regression with COHORT f.e. and AGE f.e.
Other controls: demographic variable (hhsize).
NOB: 76,900 for the more educated; 114,943 for the less educated.



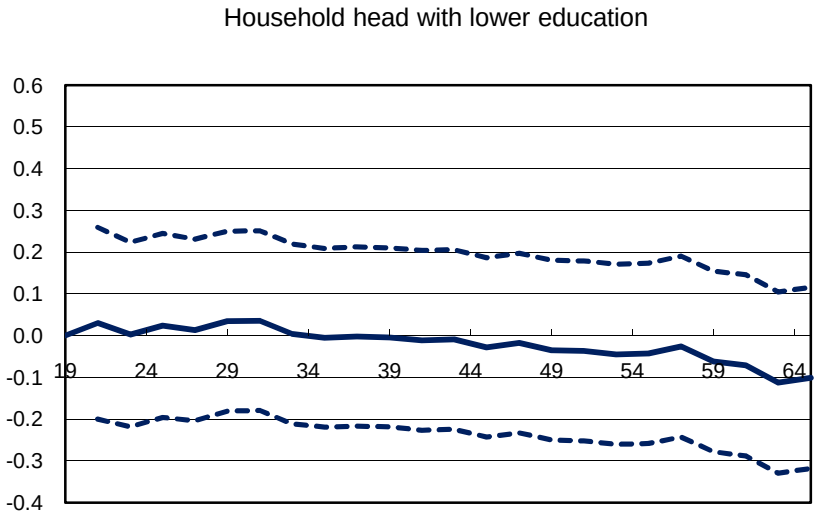
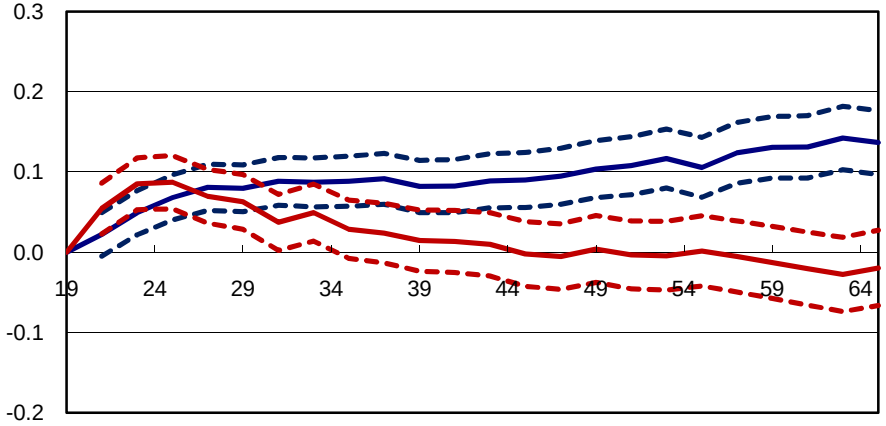


Figure 1-5.
Age effects in consumption inequality using cohort-level data, Thailand

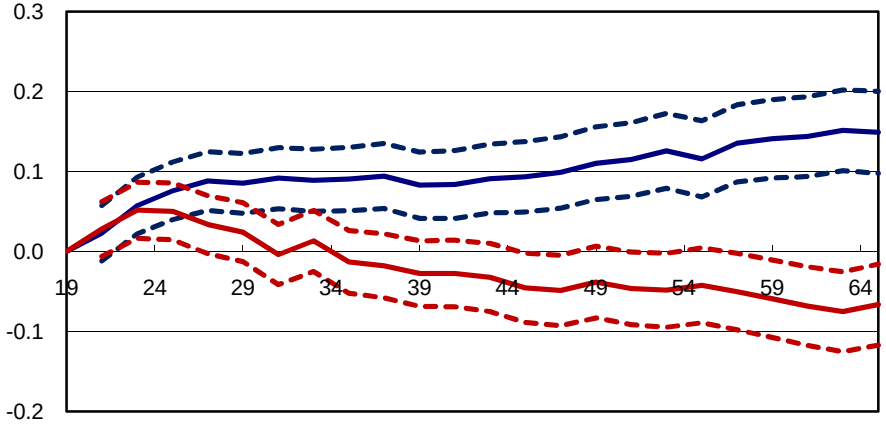
Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Coeff. on AGE f.e. (Age=19 as reference)
Source: Estimated from Thai SES data by OLS.
Cohort-level data, hh head's age in the range from 19 to 69.
Use 10 rounds from 1986 to 2004, age and cohorts in 2-year intervals.
NOB=260.

Violet: Household head with higher education
Red: Household head with lower education

Gini coefficient



General entropy with param.0



Atkinson with param.1

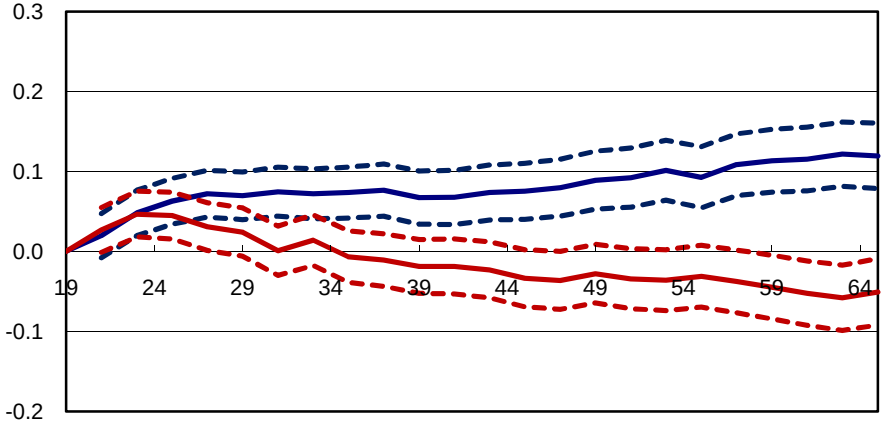
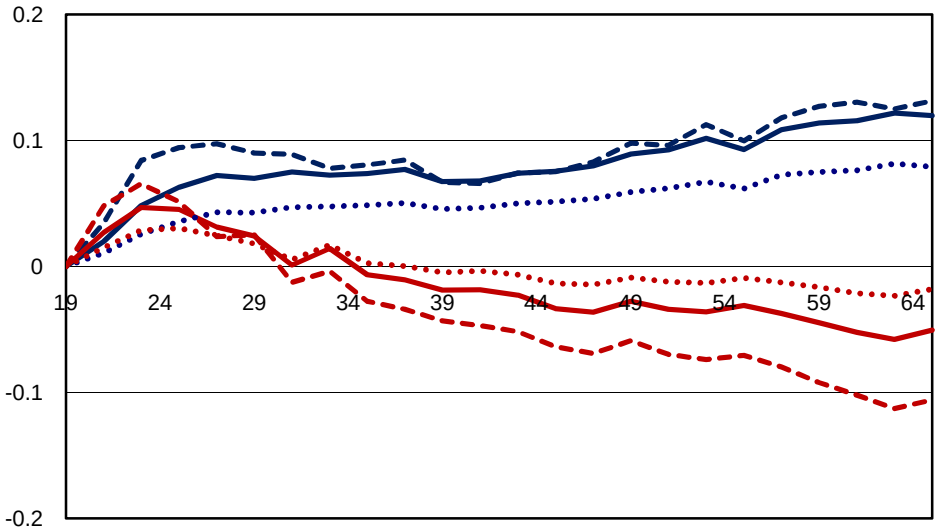


Figure 1-6.
Education contrast in age effects in consumption inequality and inequality aversion parameter, Thailand

(A) Point estimates for different inequality aversion parameters



Horizontal axis=age of the household head.
Source: Estimated from Thai SES data by OLS.
Cohort-level data, hh head's age in the range from 19 to 69.
Use 10 rounds from 1986 to 2004, age and cohorts in 2-year intervals.
NOB=260.

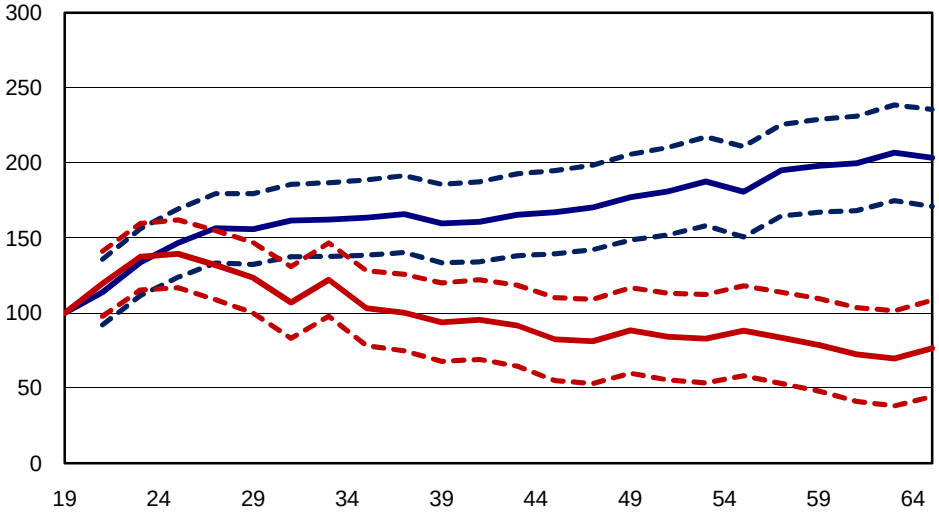
Violet: Household head with higher education
Red: Household head with lower education

(A) Inequality aversion parameter for the Atkinson inequality measures:

Vertical axis: Coeff. on AGE f.e. (Age=19 as reference)
Dot lines: parameter = 0.5 (least inequality-averse)
Usual lines: parameter = 1
Dashed lines: parameter = 2 (most inequality-averse)

(B) & (C): Dashed lines show 95% confidence interval.
Vertical axis: Index using the predicted value for Age=19 (reference) as 100

(B) Parameter = 0.5 (less inequality-averse)



(C) Parameter = 2 (more inequality-averse)

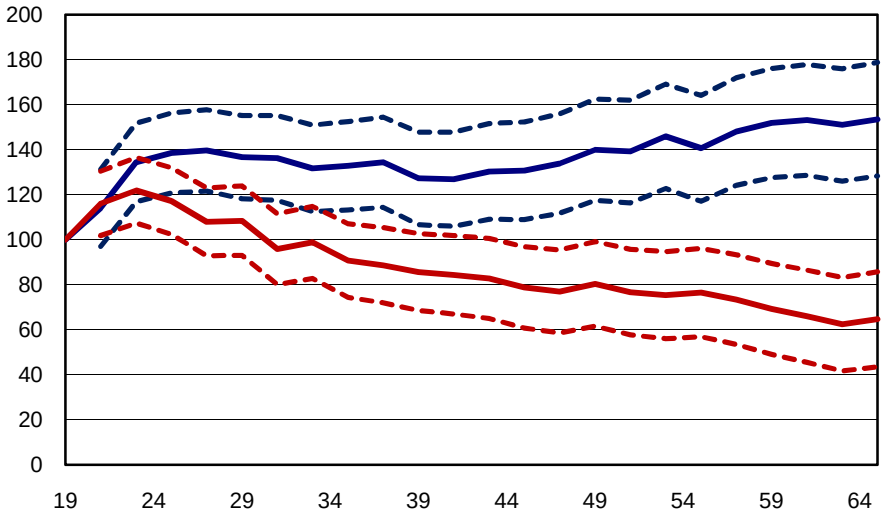
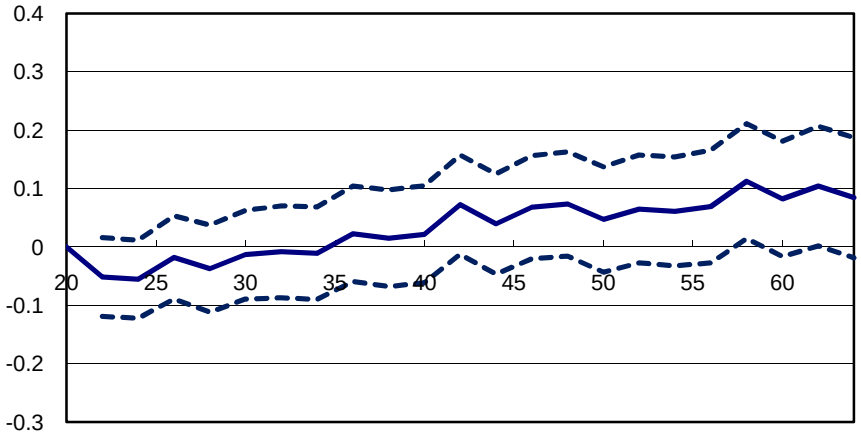


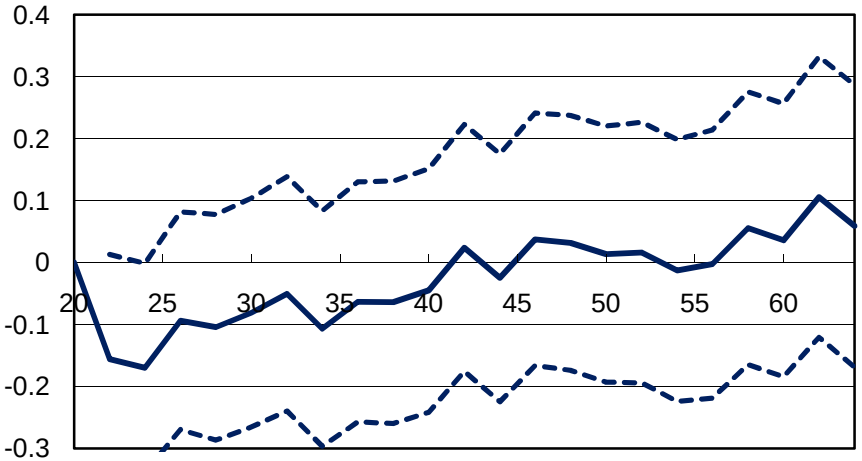
Figure 2-1.
Age effects in consumption inequality and regions, Pakistan

Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Coeff. on AGE f.e. (Age=20 as reference)
Source: Estimated from Pakistan's PIHS/PSLM data using weighted regression to control for the difference in sampling probability.
Individual-level data, hh head's age in the range from 20 to 70.
Using 4 rounds from 1998/99 to 2005/06, age and cohort f.e. in 2-year intervals
Pooled regression with COHORT f.e. and AGE f.e.
Other controls: demographic variable (hhsize).
NOB: 56,238 for all Pakistan, 21,986 for urban areas, and 34,252 for rural areas.

All Pakistan



Urban Areas



Rural Areas

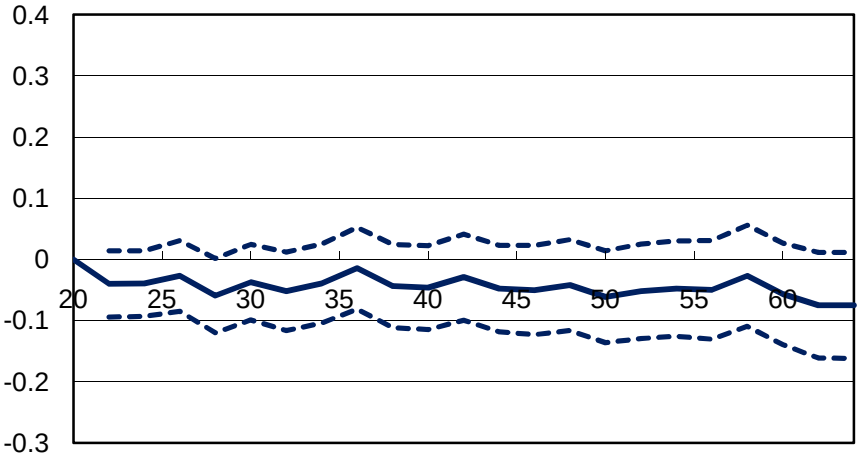
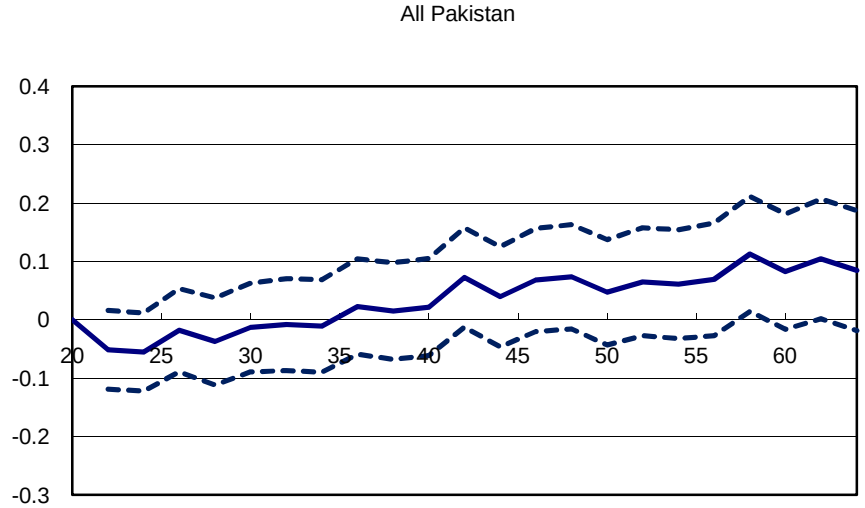
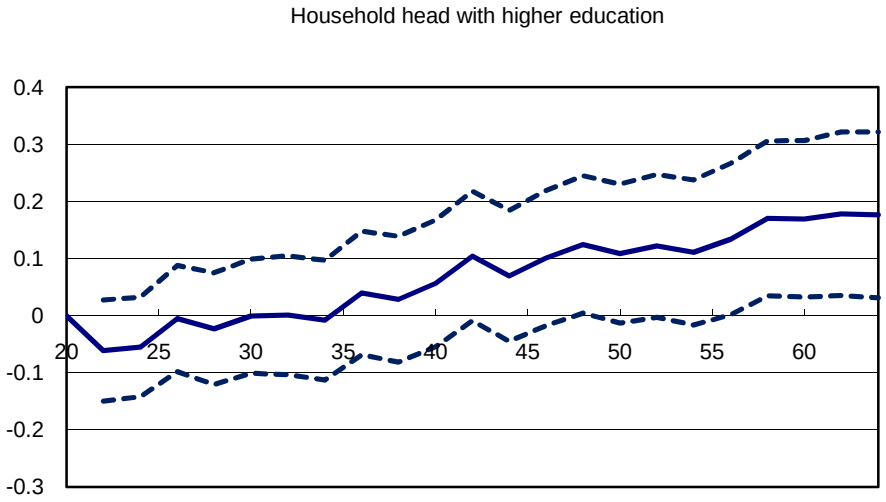


Figure 2-2.
Age effects in consumption inequality and education, Pakistan



Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Coeff. on AGE f.e. (Age=20 as reference)
Source: Estimated from Pakistan's PIHS/PSLM data using weighted regression to control for the difference in sampling probability.
Individual-level data, hh head's age in the range from 20 to 70.
Using 4 rounds from 1998/99 to 2005/06, age and cohort f.e. in 2-year intervals
Pooled regression with COHORT f.e. and AGE f.e.
Other controls: demographic variable (hhsize).
NOB: 29,628 for the more educated hhs, 26,610 for the less educated hhs.



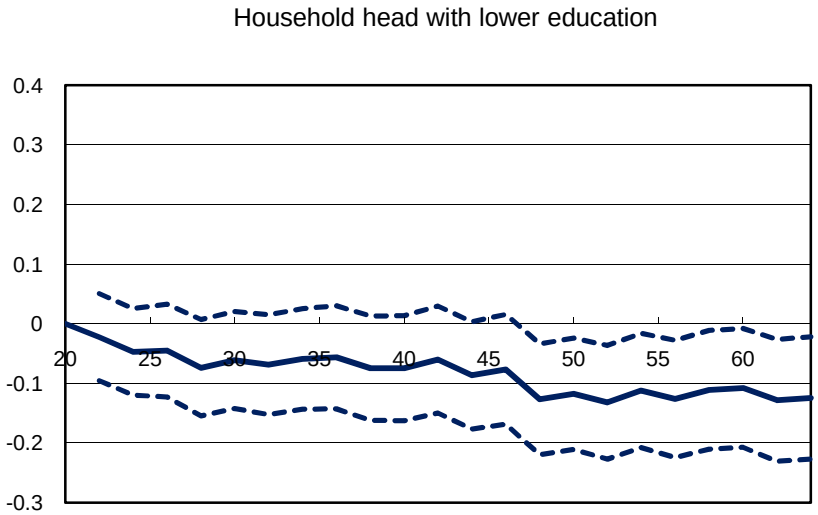
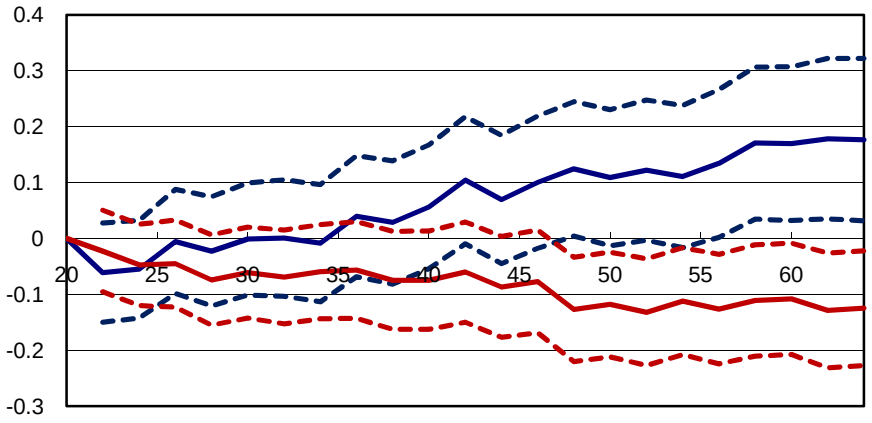


Figure 2-3.
Education contrast in age effects and household-level controls, Pakistan

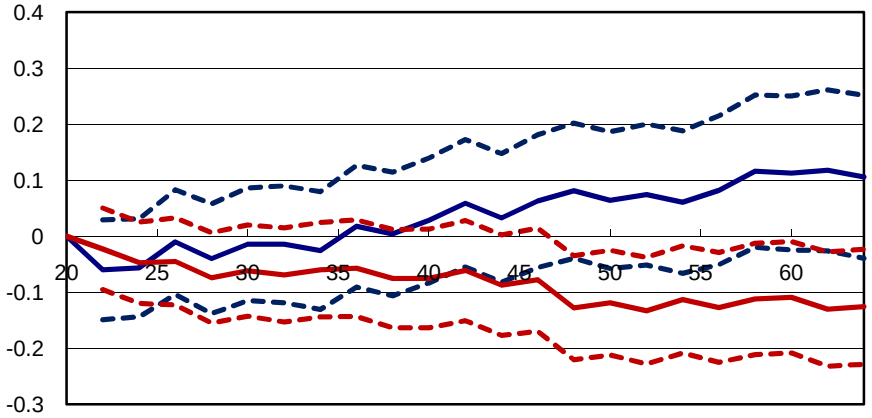
Dashed lines show 95% confidence interval.
Violet: Household head with higher education
Red: Household head with lower education

Pooled regression with COHORT f.e. and AGE f.e. and other controls:
(A) "hhsize" only (same as Figure 2-2).
(B) No household-level control
(C) Full controls of demographic and regional controls such as demographic variables (dummy for female headed hh, dummy for head's absence, dummy for missing information on head's actual birth date, size of household in numbers, and size of household in adult equivalents) and regional fixed effects (province dummies, urban dummy)

(A) "hhsize" only



(B) No household-level control



(C) Full controls of demographic and regional controls

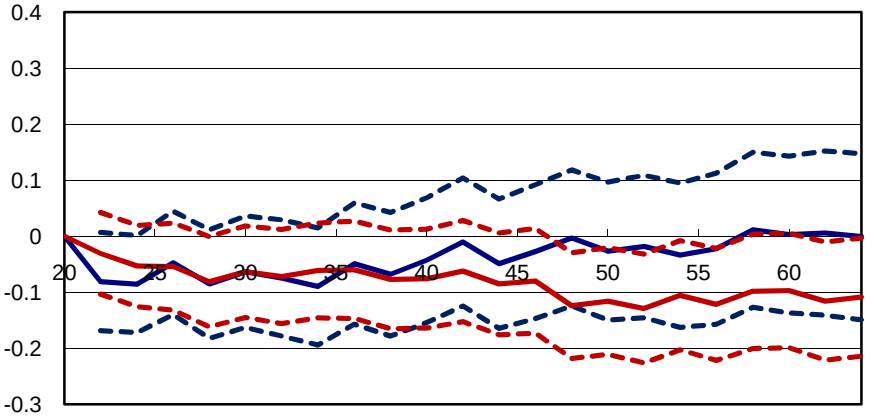
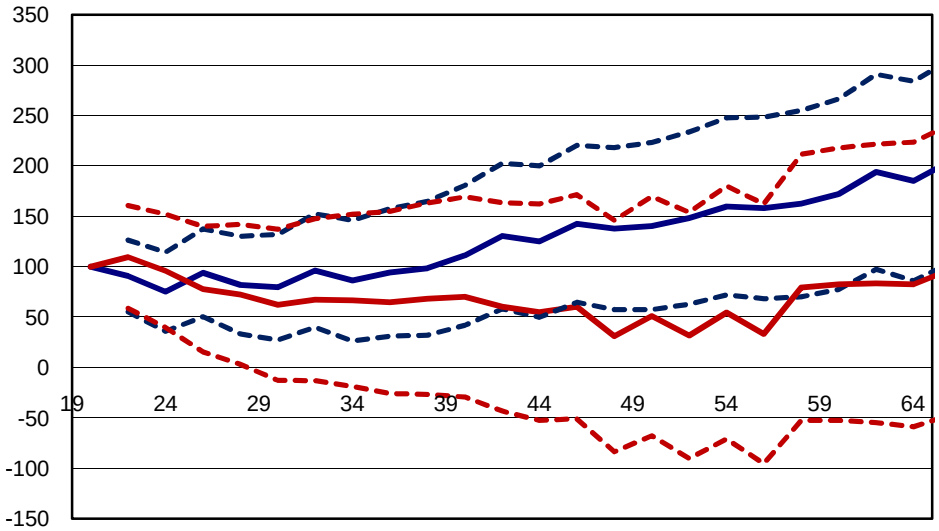


Figure 2-4. Education contrast in age effects in consumption inequality and inequality aversion parameter, Pakistan

(A) Atkinson inequality measure with parameter = 1

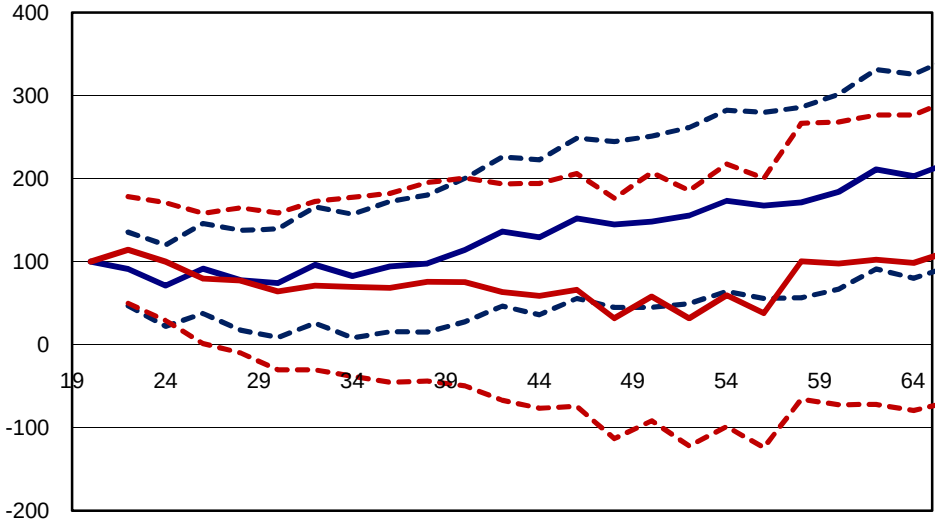


Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Index using the predicted value for Age=20 (reference) as 100
Source: Estimated from Pakistan's PIHS/PSLM data by OLS.
Cohort-level data, hh head's age in the range from 20 to 70.
Using 4 rounds from 1998/99 to 2005/06, age and cohort f.e. in 2-year intervals
NOB=104.

Violet: Household head with higher education
Red: Household head with lower education

Inequality aversion parameter for the Atkinson inequality measures is set at:
(A) Parameter = 1
(B) Parameter = 0.5 (less inequality-averse)
(C) Parameter = 2 (more inequality-averse)

(B) Parameter = 0.5 (less inequality-averse)



(C) Parameter = 2 (more inequality-averse)

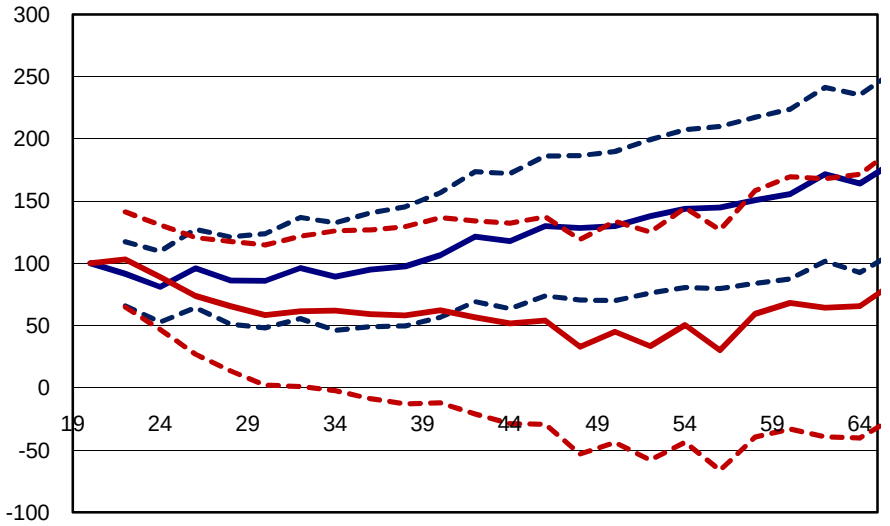
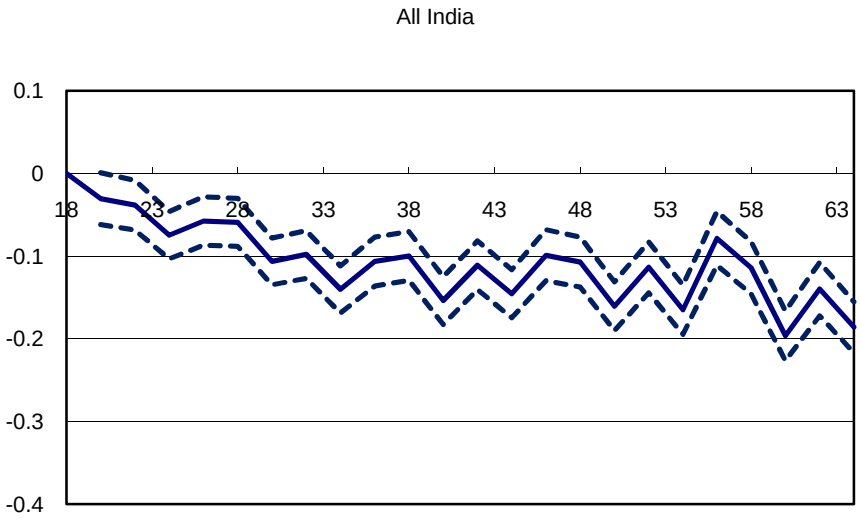
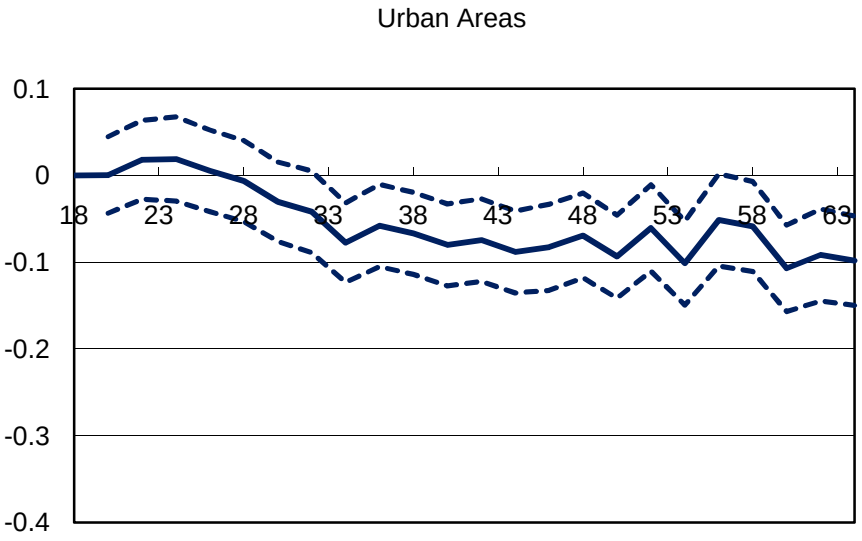


Figure 3-1.
Age effects in consumption inequality and regions, India



Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Coeff. on AGE f.e. (Age=18 as reference).
Source: Estimated from India's NSS data using weighted regression to control for the difference in sampling probability.
Individual-level data, hh head's age in the range from 18 to 70.
Use 4 rounds from 1983 to 2000, age in 2-year intervals and cohorts in 5-year intervals.
Pooled regression with COHORT f.e. and AGE f.e.
Other controls: demographic variable (hhsize).
NOB: 437,605 for all India, 162,026 for urban areas, and 275,579 for rural areas.



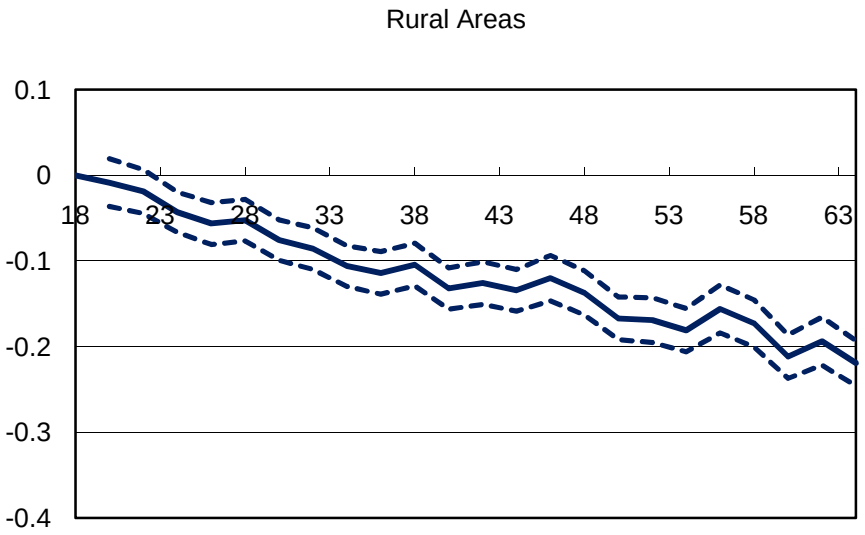
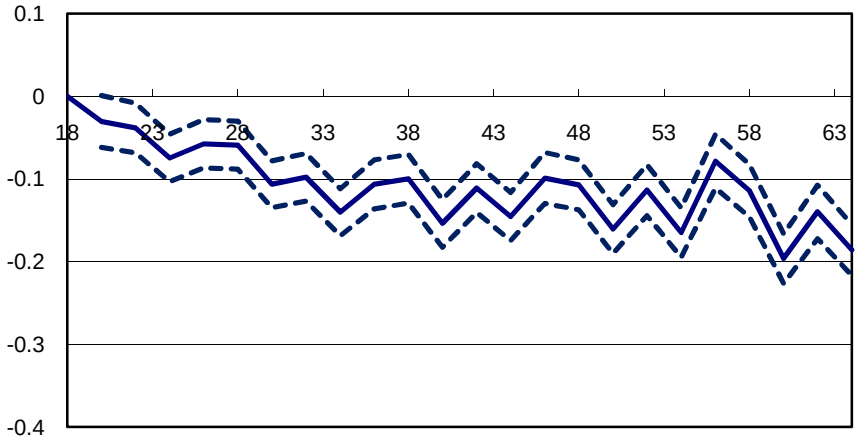
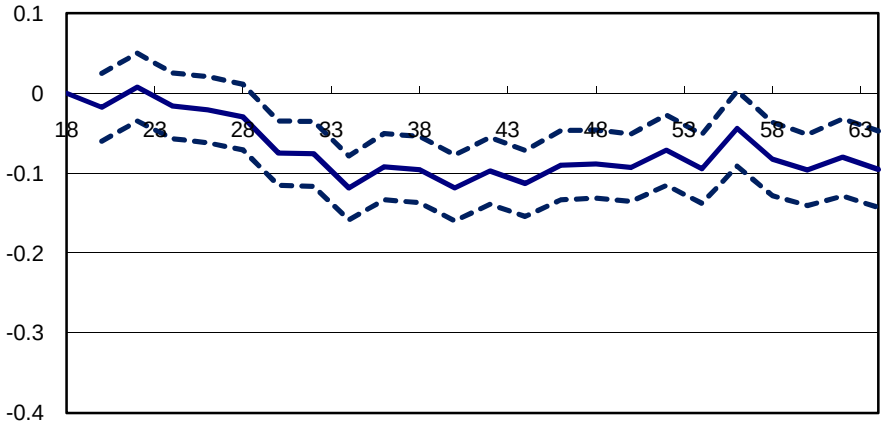


Figure 3-2.
Age effects in consumption inequality and education, India

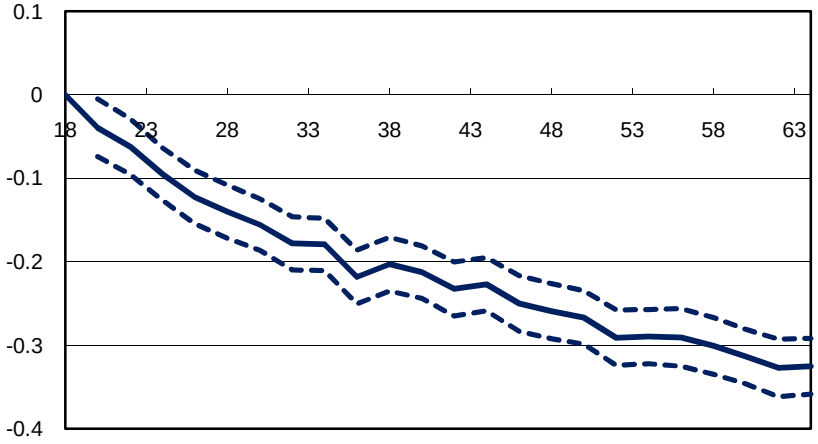
All India



Household head with higher education



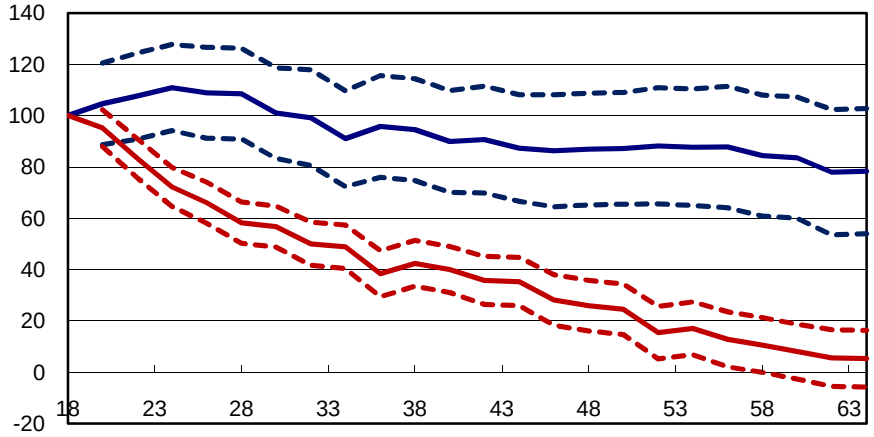
Household head with lower education



Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Coeff. on AGE f.e. (Age=18 as reference)
Source: Estimated from India's NSS data using weighted regression to control for the difference in sampling probability.
Individual-level data, hh head's age in the range from 18 to 70.
Use 4 rounds from 1983 to 2000, age in 2-year intervals and cohorts in 5-year intervals.
Pooled regression with COHORT f.e. and AGE f.e.
Other controls: demographic variable (hhsize).
NOB: 199,477 for the more educated hhs,
238,128 for the less educated hhs.

Figure 3-3. Education contrast in age effects in consumption inequality and inequality aversion parameter, India

(A) Parameter=1

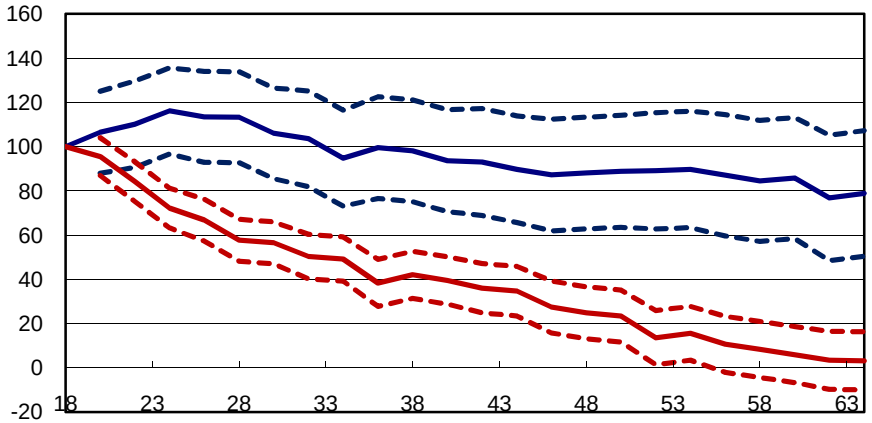


Dashed lines show 95% confidence interval.
Horizontal axis=age of the household head.
Vertical axis: Index using the predicted value for Age=18 (reference) as 100.
Source: Estimated from India's NSS data by OLS.
Cohort-level data, hh head's age in the range from 18 to 70.
Use 4 rounds from 1983 to 2000, age in 2-year intervals and
cohorts in 5-year intervals
NOB=108.

Violet: Household head with higher education
Red: Household head with lower education

Inequality aversion parameter for the Atkinson inequality measures is set at:
(A) Parameter = 1
(B) Parameter = 0.5 (less inequality-averse)
(C) Parameter = 2 (more inequality-averse)

(B) Parameter=0.5 (less inequality-averse)



(C) Parameter=2 (more inequality-averse)