UMR 8184

Thema

THéorie Économique, Modélisation et Applications

Thema Working Paper n°2011-10
Université de Cergy Pontoise, France

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January, 2011

UNIVERSITÉ de Cergy-Pontoise

CNRS CENTRE NATIONAL DE LA RECHERCHE SCIENTIFIQUE

Intergenerational earnings mobility in Japan among sons and daughters : levels and trends *

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January 28, 2011

Abstract

This paper estimates the extent of intergenerational income mobility in Japan among sons and daughters born between 1935 and 1975. Our estimates rely on a two-sample instrumental variables approach using representative data from the Japanese Social Stratification and Mobility (SSM) surveys, collected between 1965 and 2005. Father's income is predicted on the basis of a rich set of variables including education, occupation and job characteristics and we discuss changes in the Japanese earnings structure for cohorts born between the early 1900s and the 1960s. Our main results indicate that the intergenerational income elasticity (IGE) in Japan is around .3 for both sons and daughters, a rather low figure in comparative perspective. We discuss the sensitivity of the IGE to using either personal or family income as the income variable for both fathers and children. Laslty, we also examine changes across cohorts in the IGE, as well as the existence of non-linearities in the intergenerational transmission of income. Results indicate that intergenerational mobility has been roughly stable over the last decades and point to a convex relationship between parental income and child's achievement.

JEL Codes: D1, D3, J3

Keywords: Intergenerational mobility, earnings differentials, income, inequality, trends, Japan, assortative mating, education.

*We thank Jaap Dronkers, Nobuo Kanomata, Tadashi Yagi and seminar participants at Keio university and Doshisha university for useful comments. Access to the SSM data was granted through the SSM 2005 Research Committee, whose support is gratefully acknowledged. This research received financial support from the French National Research Agency, under the grant TRANSINEQ (ANR-08-JCJC-0098-01).

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1 Introduction

Over the last fifteen years, an abundant literature has analyzed the extent of intergenerational economic mobility and has revealed that in many developed countries a large fraction of economic inequality is transmitted from one generation to the next, within families (Solon 1999, Black & Devereux 2010). While available evidence has contributed to a rich description of the extent of intergenerational transmission, the mechanisms responsible for this transmission remain largely unstudied. In particular, although it has been demonstrated that more equal countries also exhibit more intergenerational mobility, as discussed for instance in Björklund & Jäntti (2009), the contribution of key ingredients of the mobility process, such as such labor market institutions, wage inequality and educational institutions is still, from an empirical perspective, unclear. To this end, the gathering of international evidence on the extent of intergenerational mobility in countries with different social and economic structure appears as an important first step. Another important limitation of existing evidence, however, in this perspective, is that it has, until recently, mostly focused on western developed countries and much less is known about the extent of intergenerational mobility in other parts of the world, including Asian countries. The objective of this paper is to fill in this gap and to measure the extent of intergenerational income mobility in Japan. Our analysis relies on the Social Stratification and Mobility (SSM) survey, a rich survey conducted between 1955 and 2005 that gathers information on individual income as well as family and social background and allows us to estimate the intergenerational income elasticity for Japan using a two-sample instrumental variables approach. We examine mobility between fathers and their sons and daughters for cohorts of children born between 1935 and 1975.

Several characteristics of Japan make it an interesting case for the study of intergenerational earnings mobility, in particular from a cross-country comparative perspective. First, Japan is often seen as a fairly equal society characterized by compressed income differentials and limited poverty. Indeed, this image seems largely sustained empirically, at least compared to other countries and until the increase in inequality that occurred in the 1990s (Gottschalk & Smeeding 2000, Tachibanaki 2009).[1] The extent to which this high degree of income and, more generally, social equality translates into a high level of economic mobility is of course an important question, both for the comprehension of contemporary Japan and for the understanding of the intergenerational income mobility process. In this respect, opposing views can be found. On the one hand, occupational and educational success are often thought, at least in Japanese

[1]Whether Japan still is, in the most recent period, an "equal" society seems highly debatable, as discussed in Tachibanaki (2009).

popular views, to depend little on family origin. On the other hand, Japan is characterized by a strongly differentiated and very competitive educational system, where success at entering the most prestigious universities, largely conditions future labor market prospects and henceforth motivates considerable financial investment by the parents. As a consequence, one would expect family background to strongly influence individual outcomes. Which view is the most relevant is the key question addressed in this paper.

Evidence on the extent of intergenerational economic mobility in Japan is currently rather limited, in comparison to the vast literature that has focused on social and educational mobility (e.g. Ishida 1993, Ojima 1998, Imada 2000, Kondo 2000). To our knowledge, only two papers have examined the extent of intergenerational income mobility in Japan.[2] Lefranc, Ojima & Yoshida (2008) provided a first assessment. As the present paper, it relies on the SSM dataset, i.e. a large representative data set covering the second half of the twentieth century. But the analysis is limited to father-son pairs and the prediction of fathers income rests on a narrow set of individual characteristics. Ueda (2009) offers estimates of intergenerational income elasticities for male and female based on the Japanese Panel Survey of Consumers. While the author tries to account for the possible biases induced, this data set suffers several weaknesses. In particular, the sample only covers the most recent period and children are observed in the early stages of their life-cycle. Parents on the contrary are observed at the same time as children which implies that for many of them, measured income corresponds to either their end of career income or their retirement income. Lastly, non-married men, who account for about 25% of the adult population are excluded from the data. All in all, these problems may influence the results and lead to unrepresentative estimates.

The main contribution of our paper is to offer a robust and in-depth analysis of intergenerational income mobility in Japan among sons and daughters. We measure mobility by the now standard intergenerational earnings elasticity (IGE), which can be obtained by regressing the log of individual annual earnings on the log of their father's earnings. In the lack of direct observation of father's income, we use a two-sample instrumental variables approach as in e.g. Björklund & Jäntti (1997). Father's income is predicted on the basis of a rich set of variables including education, occupation and job characteristics. The SSM data set used in this paper allows us to measure intergenerational mobility for a representative sample of children born between 1935 and 1975 and we also examine changes across cohorts in the IGE. Incidentally, the prediction of father's income requires us to assess changes in Japanese earnings structure

[2] A third paper is that of Yoshida (2008) but it is only available in Japanese.

for a large set of cohorts born between the early 1900s and the 1960s, using income data for the years 1965 to 1995. So an important by-product of this study is to document long-term trends in the Japanese earnings structure. We also perform several robustness analysis. First, our data includes two distinct measures of income : the first one refers to the individual's own income; the second to his or her family income. Having both measures allows in particular to isolate the contribution of marital outcomes to the intergenerational mobility process and we examine the sensitivity of the IGE to the income measure used in the analysis. Lastly, we also examine the existence of non-linearities in the intergenerational transmission process.

The main result of our analysis is that intergenerational income elasticity in Japan is around .3, which is a rather low figure in comparative perspective. This elasticity is rather similar for sons and daughters. It is constant over time for sons. For daughters, the time dynamics are a bit more complex but mobility also appears roughly stable. Lastly, in the case of sons, their overall family income seems less related to parental income than their own individual incoming, suggesting that marital sorting dampens slightly the intergenerational persistence of inequality. The rest of the paper is organized as follows. Section 2 discusses the estimation procedure and the data used in the analysis. Section 3 presents the results of the first-step estimation (section 3.1), and discusses our estimates of the extent of intergenerational mobility among sons (section 3.2) and among daughters (section 3.3). Section 4 concludes.

2 Estimation method and data

2.1 Estimation method

Most of the economic analysis of intergenerational mobility focuses on estimating the IGE in permanent (or long-term) earnings. This elasticity is given by the coefficient β in the following intergenerational regression model :

$$Y_i = \beta_0 + \beta X_i + \epsilon_i \tag{1}$$

where Y_i denotes the log of individual i's long-term earnings X_i denotes his father's long-term earnings. As abundantly discussed in the literature, β should not be seen as a structural parameter measuring the causal effect of parental resources on child's earnings, but rather as a "catch-all" descriptive measure of the intergenerational association in earnings, capturing all possible channels of transmission.

Direct estimation of equation 1 requires a considerable wealth of information. Not only does it call for a linked data set in which both father and child's earnings are observed but it furthermore requires one to observe a time-series of individual earnings in order to measure long-term earnings. Very few data sets satisfy this data requisite.

Without such data, β can be estimated using a two-sample instrumental variables (TSIV), an approach originally derived in Angrist & Krueger (1995) and Arellano & Meghir (1992). This method was first applied to the estimation of the IGE by Björklund & Jäntti (1997). The basic principle behind TSIV estimation is to replace X_i in equation 1 by a prediction $\hat{X}_i$ formed on the basis of some observable father's characteristics, Z_i.

The data requirements for TSIV estimation are significantly less stringent. The prediction is derived from a first-step equation which is estimated on a sample that is representative of the fathers' population, and in which we observe both earnings and the characteristics Z_i. Given the estimation of the first step, the data requirement for the estimation of β is to observe both child's income and father's characteristics.

TSIV has been extensively used for the estimation of the IGE and its properties are discussed in several papers including Solon (1999) and Nicoletti & Ermisch (2008). These properties depend on the choice of the instrument. If the instrument only affects child's earnings through its effect on father's earnings, the estimation of β is consistent. Indeed, in this case TSIV estimation offers the significant advantage of over-riding the *attenuation bias* that typically arises, because of classical measurement errors, when estimating equation 1 with long-term earnings replaced by current earnings (Solon 1992, Zimmerman 1992, Mazumder 2001)). However, if the instrument has a direct effect on the child's outcome, than the TSIV estimates are biased and the direction of the bias depends on the sign of the direct effect. For most of the instruments used here, the expectation is that the direct effect will be positive, hence resulting in an overestimation of the IGE. However, in practice, the order of magnitude of this overestimation turns out to be small, as discussed in Björklund & Jäntti (1997).

Another important source of bias in the estimation of the IGE is what has been referred to as the *life-cycle bias* (Jenkins 1987, Grawe 2006, Haider & Solon 2006). This bias arises when using current (usually annual) earnings instead of permanent earnings in the estimation of the IGE. In the presence of individual heterogeneity in earnings growth over the life-cycle, current earnings measures permanent earnings with error. Furthermore, it can be shown that the error is not of the classical type and is correlated with both true permanent earnings and age.[3] As a result, differences in current earnings across individuals will in general provide a

[3] The classical measurement error case refer to the situation where measurement error is independent of the

biased estimate of permanent income differentials. Since age-earnings profiles are steeper for high income individuals, current income differentials, measured at an early stage of the life-cycle, will underestimate permanent income differentials; current income at the end of the life-cycle will over-estimate permanent income differentials.

This form of measurement error will introduce an asymmetric bias in the estimation of β, depending on whether child or father's earnings are affected by this form of measurement error. Using current earnings early (resp. late) in the life-cycle, as a proxy for *child's* permanent earnings will lead to underestimate (resp. overestimate) β. Conversly, using current earnings early (resp. late) in the life-cycle, as a proxy for *father's* permanent earnings will lead to overestimate (resp. underestimate) the IGE.

To account for life-cycle biases, we follow Lee & Solon (2009) and introduce control for age and an interaction between age and father's income and use the following specification for the second-step equation :

$$Y_{it} = \alpha_t + \beta \hat{X}_i + f(age_{it}) \times \hat{X}_i + g(age_{it}) + e_{it} \tag{2}$$

where i and t are indices for individual and time. c denotes the five-year birth cohort of individual i. The α_ts denote time dummies and f and g are respectively second and fourth order polynomial functions in age.[4] $\hat{X}_i$ is predicted father's earnings at age 40; the variable *age* is normalized to zero at age 40.

Let us now turn to the first-step equation. Its purpose is to predict father's income at the age of 40 on the basis of father's education, occupation and job characteristics. In practice, we face two major issues in the specification of this equation. First, some cohorts, in particular the oldest ones, are observed far away from their mid-career. Hence, to accurately predict earnings differentials at age 40 we need to account for heterogeneity in age-earnings profiles. This is done by introducing group-specific age-earnings profiles where groups are defined on the basis of their education and employment status.[5] Second, given the large time and cohort interval used the estimation, the earnings function should allow for changes across cohorts in the earnings premium attached to the different characteristics. This is done by introducing interactions between the effect of some of the characteristics and a quadratic cohort trend.[6] In

true value.

[4]In principle, the use of polynomial function for age would allow to simultaneously include time and cohort dummies. Cohort dummies however turn out to be insignificant when added to this specification and their inclusion does not affect the results.

[5]A larger set of interaction terms could be introduced, such as occupation-age interactions. In practice, they turned out to be non-significant.

[6]The characteristics whose effects are allowed to vary across cohorts are age, education and self-employment

the end, the first-step model we estimate can be summarized by the following equation :

$$X_{ict} = \alpha_t + \phi(Z_{ic}, c) + \psi_(age_{ict}, Z_{ic}, c) + e_{ict} \qquad (3)$$

where i, c, t are indices for individual, cohort and date; Z_{ic} are the father's permanent characteristics, ϕ captures the effect of these characteristics, which is allowed to vary with c; ψ denotes the heterogenous age profile.

Lastly, predicted father's earnings is given by :

$$\hat{X}_{ict} = \hat{\phi}(Z_{ic}, c) \qquad (4)$$

2.2 Data

Our data come from the Social Stratification and Social Mobility (SSM) surveys. The SSM survey has been the primary data source for studies of social and educational mobility in Japan (Ishida 1993, Ojima 1998, Imada 2000). The first wave of the survey was conducted in 1955 by the Japanese Sociological Society. Since then, similar surveys were conducted at intervals of ten years. The earliest waves (1955, 1965 and 1975) focused only on males. A female sample was collected since the 1985 survey. The questionnaire of the last wave of the survey (2005) has also been used for similar surveys in Korea and Taiwan.

The SSM samples are designed to provide a national representative sample of the population between 20 and 70 years old. Across the different waves, the size of the male sample varies between two and three thousands individuals. The questionnaire focuses on the description of social status, educational attainment, social origin, class identification and the perception of inequality. The most important variables in our analysis are income, which is the main variable of interest, and educational and occupational attainment, which serve to predict father's income in the first-step equation. Respondents to the SSM survey are asked to report their income, education, occupation and job characteristics as well as the education, occupation and job characteristics of their father. As often the case, father's information is reported *ex post* by the survey respondent and refer to father's main occupation.

All SSM waves record two distinct measures of income. The first one is individual own income. The second one is family total income. In both cases, the variable measures annual primary income, in the year preceding the survey, before any tax or transfer and includes both labor and asset income. Furthermore, income information is available for both salaried and

status. Other interaction terms turned out to be non-significant.

self-employed workers. For most individuals of working age and who actually work, the primary component of pre-fisc income is labor earnings. Family income is the total of each family member's individual income in the respondent's household. Income is available in all waves of the survey in bracketed form, except for 1965 where income is coded continuously. The bounds and number of brackets vary across waves, between 17 and 30 main brackets. Higher income are not top-coded though and a response beyond the top category was coded continuously. In the regressions, we assign the mid-value of the bracket and use standard linear regression techniques.[7]

The education classification used in the different waves of the survey varies across waves and cohort, reflecting the changes in the Japanese educational system that occurred over the last century. For older cohorts, the classification distinguishes between five educational levels: elementary school (6 years of formal schooling), upper elementary (8 years), middle school (11 years), college (14 years) and university (17 years). For more recent cohorts, the five educational levels are: junior high school (9 years), high School (12 years), junior college (14 years), university (16 years) and graduate school (18 years). Given sample size and to assure cross-year consistency of the education classification, we used a reduced classification that distinguishes between three educational levels: lower secondary education (or lower), upper secondary education and tertiary education. This corresponds, for instance, to the classification used in Kondo (2000).

Social status is coded using the Erikson-Goldthorpe-Portocarero (EGP) classification. We use a variant of the classification in 9 groups, since a change of the questionnaire about the number of employees between 1985 and 1995 makes us impossible to distinguish between IVa (Small employer) and IVb (Independent) (Kanomata, Tanabe & Takenoshita 2008). In the end we use the following eight-categories : I- Higher-grade professionals, administrators, and officials, managers in large industrial establishments, large proprietors; II - Lower-grade professionals, administrators, and officials; higher-grade technicians, supervisors of non-manual employees; III- Routine non-manual employees, higher grade (administration and commerce) and lower grade (sales and services); IV- small proprietors and self-employed workers ; V- Lower-grade technicians supervisors of manual workers; VI- Skilled manual workers; VII- Semi- and unskilled manual workers; VIII - Farmers and farm workers. In addition to social status, SSM also includes additional characteristics of occupation and job position. We use firm size and an indicator of self-employment status.

[7]Lefranc et al. (2008) uses interval regression to deal with the bracketed form of income and show that the impact on the estimated IGE is negligible.

In the main samples used in this paper to estimate both the first- and the second-step equations, we exclude those without positive earnings in the year preceding the survey. The sample used in the estimation of the first-step equation draws on five available survey waves, 1965, 1975, 1985 and 2005. The reason for excluding 1955 is that income information is missing in that wave for farmers who accounts for a large share of the employment in older cohorts. The sample is restricted to individuals aged 30 to 59 years old. The main reason for excluding individuals older than age 60, is that many people in this sample start retiring and living on pension from this age in Japan, a problem also encountered in the study of Ueda (2009).

Second-step estimations are based on the three most recent waves (1985, 1995 and 2005).[8] For reasons already discussed, the children sample is restricted to individuals aged between 30 and 50 years old, i.e. close to the middle of their working career. For each individual in the second-step sample, we form a prediction of his father's income using estimates of the first-step equation. The prediction is based on reported father's education, EGP classifications, residential area and other occupational information. In most cases, individual in the second-step sample report their father's birth year. In this case, we use the relevant age-specific returns to education to predict father's income. When information on father's birth year is not available, we predict fathers income on the basis of the observed distribution of father's birth cohort, conditional on child's birth cohort. Mean values in the samples are presented in table 1.

3 Results

3.1 First-step estimates and long-term trends in income inequality in Japan

First-step estimates are presented in table 2 for both personal and family income. These estimates are based on the sample representative of the fathers cohorts and include only male. The first salient result that emerge from this table is the relatively limited extent of earnings differentials across the various groups. For instance, the gap betweeen the two extremes of the social ladder, higher-grade professionals and farmers, is less than .7 log points. Similarly, the difference between the low- and high-education groups is less than .2 log points. Of course, the *ceteris paribus* analysis of the coefficients is largely artificial given the collinearity of the different dimensions but even if we cumulate the above mentioned earnings gaps, they lead

[8]The reason for not using earlier waves of the SSM for the second-step is that women are only interviewed starting in 1985. Furthermore, some information on family background is missing or not recoded homogenously before 1985.

to modest earnings differentials, in comparison to what can be found in other countries. An interesting comparison can be established with the results of Lefranc & Trannoy (2005) who use similar approach and classifications. In the case of France, which is often thought to occupy an intermediate position among developed countries in terms of earnings inequality, they find a gap between the top and bottom social groups of .9 to 1 log points and of .3 to .4 between top and bottom education groups.

In contrast to occupation and education, employment characteristics lead to non-negligible earnings differentials. For instance, the benefit of being employed in a large firm is larger than the income gap between the high- and low-education groups. Self-employment status positively affects income and has a significant effect. Lastly, the main effect of the socio-demographic and employment characteristics tends to be higher in the case of personal income than in the case of family income. This result is indeed rather intuitive since we should expect other family members' income to be less correlated to the individual's characteristics than his own income. The only exception is self-employment that has a stronger impact on family than on personal income. One possible explanation is that the wives of self-employed are associated often also employed as family workers in family-owned businesses, an activity from which they would derive an income higher than the average female income.

The second part of table 2 reports the estimated age-earnings profiles by level of education and self-employment status. In the specification of the first-step equation, we only included a quadratic function of age since higher-order terms were not significant. We also present the age-earnings profile in figure 1, panel A. As expected, the slope of age-earnings profiles increases markedly with education. At the beginning of individual careers, the income effect of 10 years of individual experience amounts to about .3 log points for individuals with higher education against about .2 for individuals with secondary education or less. On the contrary, the income-age profile of self-employed individuals appears, other things equal, flatter than that of other groups.

Lastly, table 2 and figure 1, panel B, report the estimated changes across cohort in the effect of education and self-employment status. The effects reported are based on the assumption of a quadratic cohort trend. We also estimated variants of the first-step equation with piece-wise linear trends. The results were very similar. Trends across cohorts indicate a compression of earnings differentials in the long-run. First we observe a fall in the geographic income gaps. Second, the income gap between the bottom education group and the top two decreases from a high value of about .5 log points for cohorts born at the beginning of the twentieth century to a low value of about .2 for cohorts born after World War II. This reduction of the education-income gap is non-linear over the period and occurs for the most part between cohorts born in the early century and cohorts born around WWII. In contrast, earning gaps are rather constant across cohorts born in the 1945-1965 interval. This time pattern roughly coincides with the pace of educational expansion that occurred in Japan throughout the twentieth century. As shown in figure 2, educational attainment rised markedly over this period, with the bulk of the educational expansion occurring between the early century cohorts and the cohorts born in the late 1940s. At the end of our period, the downward trend in earnings differentials seems inverted and we observed a slight increase in earnings inequality.[9] This small rise in inequality is however much smaller than the one documented in Tachibanaki (2009). It may be due in particular to the fact that our first-step equation is estimated on a sample of males aged between 30 and 59 and born before 1975, who were less affected by the rise in earnings inequality.

3.2 Intergenerational income mobility for sons

Main results Our main estimates of intergenerational economic mobility for sons in Japan are given in columns (1) to (3) of table 3. Altogether, they suggest a value of the IGE around .35, although the value varies slightly with the variables used to measure children and fathers incomes. Most values reported in international assessment of the intergenerational mobility are estimated using personal labor income and are thus largely comparable to our estimate in column (1), which is based on reports of both father's and son's personal income. The major discrepancy in the measurement of income, between the data used in column (1) and most international estimates is the inclusion of asset income in our measure of personal income. Strictly speaking, estimates in column (1) represent an average of the intergenerational elasticity for labor earnings and for asset income, weighted by the contribution of both sources to personal income. Previous estimates indicate that the intergenerational elasticity for wealth is significantly larger than for labor earnings (e.g. Mulligan 1997). So for the purpose of international comparisons of the extent of the mobility in the personal labor earnings, the figure of .33, in column (1), should be seen as an upper bound estimate of the elasticity in Japan. Of course, the order of over-estimation is probably rather small, given the share of asset income in overall income.

Compared to column (1), column (2) regresses sons' personal income on the overall family income of their parents. Measuring parental economic status by means of the total family

[9] One may suspect that this small positive trend is artificially induced by the choice of a quadratic specification for cohort effects. In fact, it is clearly not the case and the specification of piece-wise linear trends with a break around cohort 1945 also indicate an increase of education income differentials at the end of the period.

income, instead of the sole father's income leads to a slightly higher IGE of .37 (instead of .33). This suggests that total family income may better capture the broad nexus of factors that shape individual success. This confirms results already obtained in several other papers. In fact, in Japan, the gap between the two estimates appears smaller than what has been found in other countries, for instance the United States. For this country, Solon (1992) indicates a difference in the estimated IGE using both measures of .09, with the IGEs being respectively .39 and .48 when using father's earnings or family income.[10]

Lastly column (3) provides estimates of the IGE based on the regression of sons' family income on their parents' family income. Compared with estimated in column (2), since family income aggregates the income of both spouses, estimates based on this variable will reflect both the influence of assortative mating on mobility, and the direct intergenerational transmission between parents and children.[11] More precisely, as discussed in Chadwick & Solon (2002), the estimated IGE in column (3) is the weighted average of (i) the intergenerational elasticity of son's own income to his parents' family income and (ii) the elasticity of his spouse's personal income to the son's parents' income, where the weights are given by the share of each of the two spouses' income in the total family income. Using family income (column 3) rather than personal income (column 2) as a dependant variable reduces the IGE from .37 to .30. This suggests a relatively low elasticity between the wife's own income and her parents-in-law family resources. Ermisch, Francesconi & Siedler (2006) report similar results for the US.

Comparison with previous estimates Overall, our results point to a value for the IGE for sons in Japan slightly higher than .3. This represents an intermediate degree of intergenerational mobility, compared to other developed countries. The results found here also differ slightly from previous results obtained for Japan. First, the IGE reported in table 3 are slightly higher than those reported in our previous study (Lefranc et al. 2008), which largely excluded self-employed workers from the analysis and suggested a value of the IGE in the inteval 0.22-0.31. Second our results are also significantly lower than those of Ueda (2009) who reports an IGE for men in the interval 0.41–0.46. Several differences in the sample selection procedure are likely to account for this discrepancy. First, for males, the results of Ueda only concern married individuals, since single men, who represent about 30% of the relevant population, are excluded from the *Japanese*

[10]Other evidence on the incidence of the measure of family economic status in the US can be found in Altonji & Dunn (1991) and Mulligan (1997). Our findings of a smaller gap between the IGEs find using both variables may be partly attributable to the fact that our measure of father's income, contrary to earnings measures used in the US, already incorporates the asset earnings of the father.

[11]For unmarried individuals, family income will typically coincides with individual income, except in the rare case of individuals living with parents or relatives.

Panel Survey of Consumers that she uses.[12] If there is selection into the marriage market this will bias her estimates, presumably downward though. Second, income is observed at the same time for children and parents. Hence children are observed fairly early in their life-cycle and parents are observed fairly late. Ueda accounts for possible life-cycle biases by controlling for age effects but this strategy is only valid if the earnings structure has remained constant over time. On the contrary, section 3.1 indicates that earnings differentials have diminished over time. So other things equal, this implies that end of life-cycle observations may lead to underestimate inequality among parents and consequently overestimate the IGE. The fact that some parents have already started collecting a pension will also presumably go in the same direction of underestimating inequality among parents and the overstating the IGE.

Sensitivity analysis We performed two main tests of sensitivity of our results to changes in the specification. First we examined the possibility of changes over time in the value of the IGE. To this end, we interact our measure of predicted father's income with a dummy variable equal one for cohorts born after 1952. The choice of the cut-off date roughly splits our sample in two equal-sized groups and allows to separate the most recent cohorts who entered the labor market after the period of high economic expansion that followed World War II. Results are given in columns (4) to (6). They indicate that the IGE was remarkably stable across cohorts over the period studied here. This converges with the results of Ishida & Miwa (2008) who find that social mobility has remained stable in Japan in the second half of the twentieth century.

Second, we also tested for the presence of non-linearities in the relationship between father and son's income. This is done by regressing child's income on a polynomial function of the log of father's income. Columns (7) to (9) report the estimates obtained when using a quadratic function of father's income, for the different combination of measures of child's and father's income. We also tried using cubic and quartic expansions of father's income. Results are given in figure 3. When focusing on son's own income, we do not find any compelling evidence of a non-linear relationship. Higher-order terms of father's predicted income never turn out to be significant and the predicted relationship always lies very close to the linear prediction.[13] However, a convex relationship appears when using family income as the explained variable, which points to a non-linearity in the effect of assortative mating, the association between son's parental income and wife's income being higher at the top of the income distribution than at

[12] According to the 2005 census, 30% of men between 30 year old and 59 year old and 25% of women are single.

[13] This is similar to the results reported for the US and UK by Bratsberg, Røed, Raaum, Naylor, Jäntti, Eriksson & Österbacka (2007), although the same study also documents a convex relationship in Nordic country that may partly account for the low overall IGE in Scandinavia.

the bottom. We now turn to the analysis of intergenerational transmission for daughters.

3.3 Intergenerational income mobility for daughters

Main results Our main estimates of the intergenerational elasticity for daughters are given in table 4, columns (1) to (5). OLS estimates of the IGE for own income appear particularly low, with a value below .2. One should however note that personal income is missing for about 40% of our daughters sample. This may in particular result from non-participation to the labor market. If the non-report of personal income is related to parental income, estimates in column (1) are unrepresentative. Thus, in column (2), we re-estimate the IGE using Heckman's sample selection model. In the selection equation, an indicator for the report of income is regressed on marital status, indicators of the number of children, spouse's income and education when married, as well as age profile and year indicators. Overall, the results in column (2) indicate negative sample selection : women with the greatest income potential are less likely to report their own income, because of lower participation to the labor market, which leads OLS to underestimate the extent of the IGE. Once sample selection is taken into consideration, the IGE rises to a value of slightly lower than .3, which is indeed very comparable to the one found for sons.

Columns (4) and (5) report the same exercise for the combination of daughter's family income and parental family income. In this case, taking sample selection into account does not affect the measured IGE. In the sequel, we only report results derived from the sample selection model.

As in the case of sons, the IGE appears slightly higher when regressing child's income on total parental income rather than father's personal income. However, contrary to what we find for men, the IGE is not sensitive to whether daughter's income is measured by personal income (.313 in column (3)) or by family income (.314 in column (5)). In the light of the model developed by Chadwick & Solon (2002), this first suggests a rather high correlation between daughter's parents' income and her husband's own income. Compared to what we found for sons, this points to a gender asymmetry in the elasticity of own income to parents-in-law's income. This asymmetry could largely be explained by a more compressed earnings distribution for daughter's in Japan, together with gender neutral correlation in the characteristics of spouses and their in-laws.

Sensitivity analysis As in the case of sons, we also check for the presence of changes across cohorts in the intergenerational transmission of earnings as well as non-linearities in the transmission process. Results for the presence of non-linearities are given in columns (9) to (11). All in all, squared-income and higher order terms are never significant. The graphs presented in figure 3, panels A and B, go in the same direction and suggest that the relationship between parental income and adult income - both personal and family income - can be adequately summarized by a linear relationship.

Results for changes across cohorts are given in colums (6) to (8), in table 4. At first sight, they indicate opposing results, depending on the variable one uses to measure daughter's income. When using daughter's personal income, as in column (6), the striking result that emerges is not only that the IGE for recent cohorts is markedly higher for cohorts born after 1952, but, most importantly, that the value of the IGE for daughters born before 1952 is not significantly different from zero. Consequently, the results in column (6) *prima facie* suggest a drastic switch from a regime where the intergenerational transmission of inequality was by and large absent, contrary to what was observed for sons, to a new situation where intergenerational mobility is substantially lower, and in fact lower than for sons.

This image is largely misleading and results from the specificities of the female labor market in Japan. For women born before 1952, employment concentrates in low-status and low-wage occupations, so called 'pink-collar' jobs, where the status attainment process is very limited compared with men (Imada 1998) with little promotion perspectives (Suzuki 2005). A large share of part-time work is also observed, together with a high share of family employment. This leads to a fairly compressed wage structure that will be reflected directly into a low IGE. But the primary meaning of this low IGE is that labor market status in general, and own income in particular, represents a poor proxy of social status and mobility for women in the older cohort, given in particular the small share of women's personal income in the overall family income.[14] At the same time, for that same cohort, family background is very highly correlated to the family income earned in adulthood as well as with several over dimensions of individual attainment, such a education. And column 8 indicates that the intergenerational elasticity for family income essentially stayed unchanged across cohorts, and, if anything, has declined slightly for more recent cohorts.

4 Conclusion

In this paper, we have examined the extent of intergenerational transmission of income inequality for sons and daughters in Japan. Our estimates suggests that a value of the IGE around

[14]For our two cohort groups, this share amounts respectively to an average of .2 and .22.

.3 adequately summarizes the persistence of inequality between fathers and children for both sons and daughters. Overall, this value puts Japan in an intermediate position, compared to other developed countries, on the scale of intergenerational mobility. For instance, among the countries surveyed in Björklund & Jäntti (2009), Japan appears more mobile than the United States, the United Kingdom, Italy and France, all of which exhibit an IGE greater than .4, comparable to Germany but less mobile than Scandinavian countries, Australia and Canada.

At first sight, this rather high degree of mobility may seem hard to reconcile with the alleged importance of parental investment in child's education at work in Japan. As already abundantly described in several studies, access to higher education in Japan is often expensive and selective, forcing family to elaborate complex educational strategies and to undertake significant financial investments to support them. For instance families cover between 71 and 86 % of the annual expenditures of university students (Kondo 2000). Furthermore, besides tuition fees, parents often invest significant amounts in "shadow education" such as cram schools and private tutoring.

Two factors, however are likely to limit the incidence of these investments for the transmission of income across generations. Both factors, emphasize the role of compressed earnings differentials, but at different stages of the intergenerational mobility process. The first argument emphasizes the small degree of inequality among Japanese parents. In such a context, even if parents invest a large share of their income in their children's education, in the end, the resulting distribution of human capital in the next generation will also be relatively equal. This is confirmed by the analysis of "shadow education" undertaken in Stevenson & Baker (1992), who emphasize the following three aspects of private investment in education in Japan. First, family financial investment is high on average. Second, such investment is efficient at improving educational attainment. But third, financial investment and the use of shadow education seem to vary little with characteristics of the family background such as parental education or family income. The second argument applies to the end of the intergenerational transmission process. As discussed for instance in Solon (2004), lower returns to human capital, as seems to be the case in Japan compared to most developed countries, will translate into a lower IGE and limit the income consequences of inequalities of parental investment.

Part of the high degree of intergenerational mobility observed in our sample may also be explained by the specificities of the high economic expansion period that develop after World War II. Of course, the high aggregate growth that Japan experienced in this period came along with a wave of rapid industrial development and occupational change that may have fostered intergenerational mobility. However, it is important to stress that our results do not suggest any slowdown in the mobility process for the cohorts of children born after 1952 and who entered the labor market after the high-growth era. Whether this relatively high degree of mobility will be maintained in the face of the recent rise in earnings inequality described in for instance in Tachibanaki (2009) is of course an open question.

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Table 1: Descriptive statistics

		Fathers sample		Sons sample		Daughters sample	
		mean	sd	mean	sd	mean	sd
birth year		1941.127	14.957	1953.814	9.901	1956.203	9.702
survey year	1965	0.178		0.000		0.000	
	1975	0.230		0.000		0.000	
	1985	0.210		0.371		0.224	
	1995	0.197		0.336		0.411	
	2005	0.184		0.293		0.365	
social status (EGP class)	I- Higher service	0.118		0.145		0.028	
	II- Lower service	0.183		0.229		0.181	
	III- Routine clerical/sales	0.105		0.130		0.428	
	IV- Small employer and Independent	0.167		0.134		0.050	
	V- Manual foreman	0.081		0.095		0.005	
	VI- Skilled manual	0.100		0.085		0.073	
	VII- Semi-Unskilled manual	0.153		0.149		0.220	
	VII- Farmers/Farm workers	0.094		0.033		0.015	
education	primary or lower secondary education	0.344		0.129		0.099	
	upper secondary education	0.419		0.501		0.601	
	higher education	0.237		0.370		0.300	
firmsize	>1000	0.164		0.187		0.099	
	government	0.106		0.119		0.091	
	other	0.730		0.694		0.810	
self-employed		0.321		0.236		0.196	
residential area	6 major cities	0.137		0.125		0.134	
	urban	0.584		0.643		0.641	
	rural	0.279		0.232		0.225	
income	ln(individual income)	5.940	0.687	6.190	0.531	5.003	0.913
	ln(family income)	6.203	0.650	6.449	0.518	6.495	0.560
N		7170		2656		2598	

Notes: sd refers to the standard deviation; N refers to the maximum number of observation. Actual numbers may vary because of missing observations. Fathers and sons samples exclude individuals with missing or zero earnings. Individuals with zero or missing earnings are kept in the descriptive statistics for female except for computation of mean income.

Table 2: First-step estimates

	(1)		(2)	
	ln(individual income)		ln(family income)	
EGP class I	REF		REF	
EGP class II	-0.205***	(0.0239)	-0.149***	(0.0234)
EGP class III	-0.430***	(0.0275)	-0.302***	(0.0270)
EGP class IV	-0.422***	(0.0285)	-0.394***	(0.0277)
EGP class V	-0.292***	(0.0301)	-0.221***	(0.0295)
EGP class VI	-0.522***	(0.0290)	-0.408***	(0.0285)
EGP class VII	-0.545***	(0.0268)	-0.422***	(0.0264)
EGP class VIII	-0.680***	(0.0326)	-0.571***	(0.0318)
primary or lower sec. education	REF		REF	
upper secondary education	0.134***	(0.0238)	0.135***	(0.0233)
higher education	0.185***	(0.0293)	0.174***	(0.0287)
firm size : >1000	0.263***	(0.0177)	0.198***	(0.0172)
firm size : government	0.131***	(0.0216)	0.128***	(0.0211)
self-employed	0.174***	(0.0293)	0.262***	(0.0286)
residential area : 6 major cities	REF		REF	
residential area : urban	-0.0769***	(0.0235)	-0.0221	(0.0228)
residential area : rural	-0.117***	(0.0264)	0.00367	(0.0257)
age $\times$ prim./low.sec. educ.	0.0135***	(0.00297)	0.0111***	(0.00303)
age $\times$ upper secondary educ.	0.0156***	(0.00287)	0.0126***	(0.00290)
age $\times$ higher educ.	0.0225***	(0.00322)	0.0146***	(0.00322)
age $\times$ self-employed	-0.000332	(0.00255)	0.0000385	(0.00249)
age^2 $\times$ prim./low.sec. educ.	-0.000414*	(0.000231)	-0.000139	(0.000228)
age^2 $\times$ upper secondary educ.	-0.000513***	(0.000190)	-0.000290	(0.000191)
age^2 $\times$ higher educ.	-0.000564**	(0.000229)	-0.000277	(0.000229)
age^2 $\times$ self-employed	-0.000445**	(0.000207)	-0.000361*	(0.000202)
cohort $\times$ age	0.0000983	(0.000189)	-0.0000382	(0.000198)
cohort $\times$ age^2	0.00000894	(0.0000110)	-0.00000493	(0.0000110)
cohort $\times$ upper secondary educ.	-0.00227	(0.00159)	-0.00399**	(0.00173)
cohort $\times$ higher educ.	-0.00190	(0.00173)	-0.00190	(0.00186)
cohort $\times$ self-employed	-0.00401***	(0.00119)	-0.00266**	(0.00120)
cohort $\times$ residential area : urban	0.00291**	(0.00134)	0.00251*	(0.00134)
cohort $\times$ residential area : rural	0.00607***	(0.00154)	0.00736***	(0.00158)
cohort^2 $\times$ age	0.00000834	(0.00000553)	0.00000853	(0.00000565)
cohort^2 $\times$ age^2	-0.000000460	(0.000000363)	-0.000000249	(0.000000368)
cohort^2 $\times$ upper secondary educ.	0.000149**	(0.0000679)	0.0000527	(0.0000726)
cohort^2 $\times$ higher educ.	0.000179**	(0.0000784)	0.000175**	(0.0000825)
cohort^2 $\times$ self-employed	-0.0000867	(0.0000547)	-0.0000434	(0.0000548)
cohort^2 $\times$ residential area : urban	-0.0000717	(0.0000672)	-0.000102	(0.0000671)
cohort^2 $\times$ residential area : rural	-0.000160**	(0.0000748)	-0.000192**	(0.0000757)
year 1975	0.466***	(0.0274)	0.475***	(0.0266)
year 1985	0.584***	(0.0379)	0.630***	(0.0372)
year 1995	0.708***	(0.0507)	0.834***	(0.0514)
year 2005	0.569***	(0.0799)	0.697***	(0.0847)
Constant	5.731***	(0.0394)	5.764***	(0.0384)
Observations	7170		6814	
R-squared	0.459		0.452	

Standard errors in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Notes: age is centered at age 40; cohort is centered at year 1945.

Table 3: IGE estimates - sons

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
ln (father's income)	0.336***	0.374***	0.308***	0.341***	0.375***	0.306***	-0.464	-1.160	-1.950**
	(0.0424)	(0.0485)	(0.0484)	(0.0565)	(0.0640)	(0.0628)	(0.670)	(0.803)	(0.800)
ln (father's income)				-0.0170	-0.00697	0.0119			
× cohort born after 1952				(0.0865)	(0.0996)	(0.100)			
ln (father's income)^2							0.0723	0.134*	0.198***
							(0.0604)	(0.0703)	(0.0700)
type of child's income	own	own	family	own	own	family	own	own	family
type of father's income	own	family	family	own	family	family	own	family	family
N	2273	2273	2113	2273	2273	2113	2273	2273	2113
R-sq	0.149	0.146	0.135	0.151	0.148	0.137	0.149	0.147	0.138

Notes: Standard errors in parentheses. * p<0.10, ** p<0.05, *** p<0.01. Columns (1) to (9) report IGE estimates for sons based on equation 2 for different combinations of father's and child's income variables.

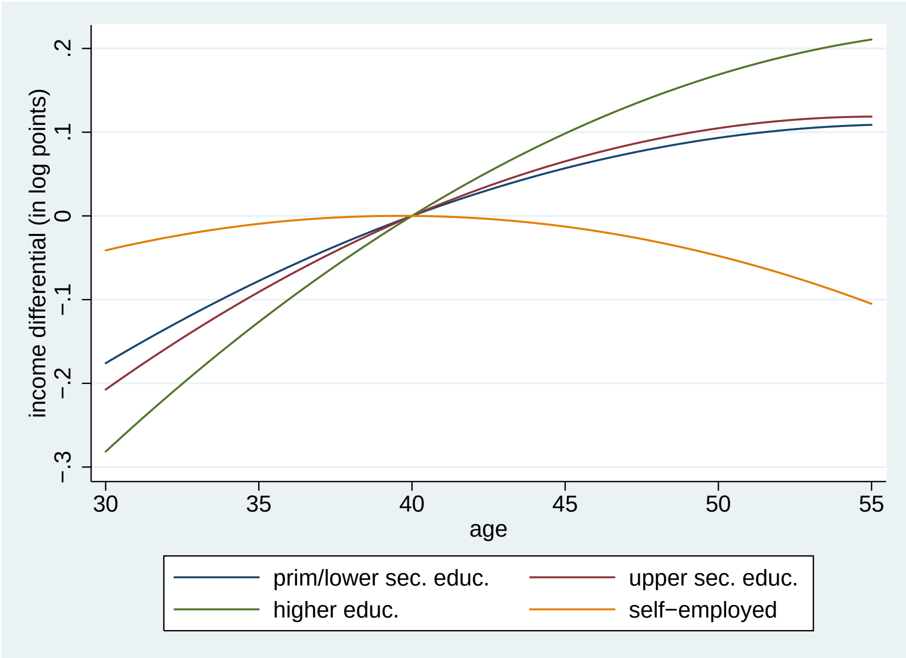
Table 4: IGE estimates - daughters

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
ln (father's income)	0.182**	0.280***	0.313**	0.328***	0.314***	-0.00715	0.103	0.367***	-0.935	-2.006	1.034
	(0.0903)	(0.0922)	(0.122)	(0.0576)	(0.0572)	(0.158)	(0.193)	(0.0909)	(1.348)	(2.637)	(1.351)
ln (father's income)						0.438**	0.348	-0.0829			
$\times$ cohort born after 1952						(0.193)	(0.248)	(0.118)			
ln (father's income)^2									0.121	0.198	-0.0614
									(0.134)	(0.225)	(0.115)
type of child's income	own	own	own	family	family	own	own	family	own	own	family
type of father's income	own	own	family	family	family	own	family	family	own	family	family
Model	OLS	heckman	heckman	OLS	heckman	heckman	heckman	heckman	heckman	heckman	heckman
ρ		-0.799	-0.796		-0.620	-0.810	-0.810	-0.618	-0.800	-0.798	-0.620
N	1561	2462	2462	1927	2462	2462	2462	2462	2462	2462	2462

Notes: Standard errors in parentheses. * p<0.10, ** p<0.05, *** p<0.01. Columns (1) to (11) report IGE estimates for daughters based on equation 2 for different combinations of father's and child's income variables. Columns (1) and (4) are estimated using Ordinary Least Squares. Other columns are estimated using the Heckman's sample selection model. The selection equation uses the following regressors : marital status, indicators of the number of children, spouse's income and education when married, age profile and year dummies.

Figure 1: Age and cohort effects, first-step income equation - personal income

A- Age-earnings profiles



B- Cohort effects

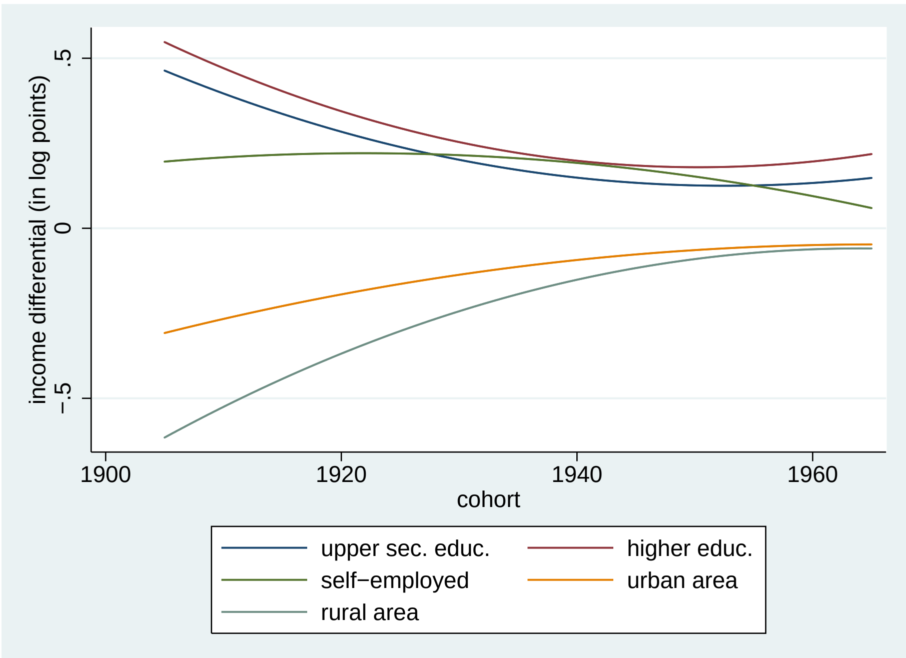


Figure 2: Distribution of education by birth cohort

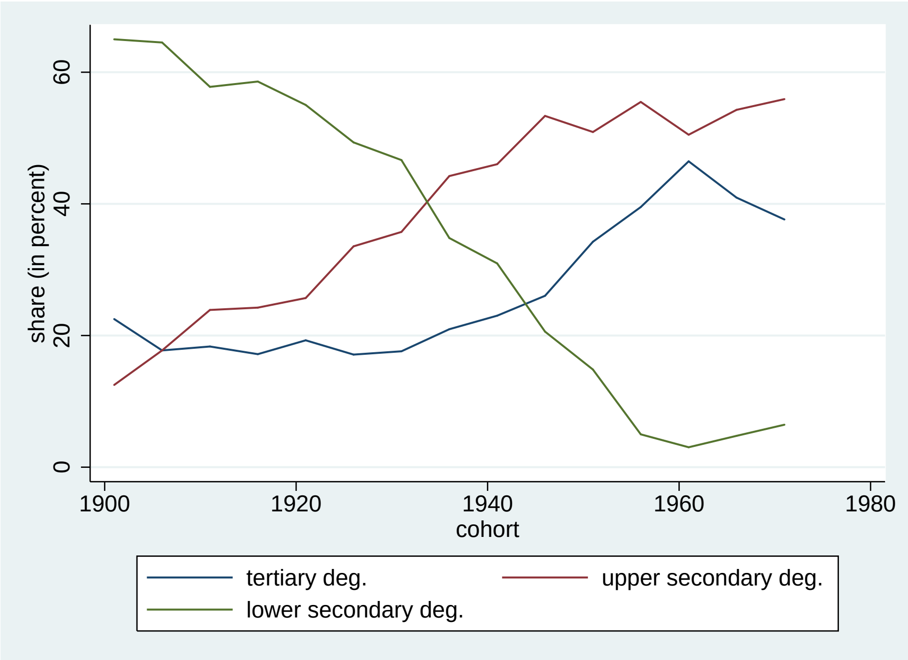


Figure 3: Non-linearities in the intergenerational transmission of income

A- son's own income on father's own income

B- son's family income on father's family income