Child Poverty, Generational Mobility and the One Child Policy in Urban China

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Abstract.

Chinas' controversial and far reaching One Child Policy (OCP) introduced in 1979 changed fundamentally the nature of both existing and anticipated marriage arrangements and influenced family formation decisions in many dimensions especially with respect to the number of and investment in children. It may well be expected to have influenced their well being and life chances. Child poverty, or the over-representation of children in the poverty group, has been a major policy issue in western societies. In the United Kingdom and Canada its elimination has been a declared policy target, in the USA, its deleterious consequences have been attacked with policies promoting generational mobility (i.e. policies to reduce the dependence of child outcomes on parental circumstances) under an equal opportunities imperative. Here the impact of the OCP on child poverty and generational mobility in the context of the parent/child educational attainments and incomes is considered. Using data drawn from an urban household survey carried out in six provinces in China, namely Shaanxi, Jilin, Hubei, Sichuan, Guangdong and Shandong (the first two may be considered low, the second two intermediate and the third pair high income provinces) the impact of the OCP on child poverty and generational mobility is studied.

The extent to which the OCP influenced investment in children is studied by studying the way in which the relationship between the educational attainment of children and family characteristics changed with the introduction of the OCP. Broadly speaking the impact of household income and parental educational attainment increased significantly over time, there also appears to be an emerging negative household size effect and a negative birth order effect suggesting that the level of investment in children diminished with the size of the family. Finally a positive gender effect emerged (girls advanced more than boys). Examination of the proportion of children in the poverty group (defined by incomes below various proportions of median income) revealed that, unlike western societies, children are not overrepresented in the poverty group, neither before nor after the OCP. Here comparisons are made with Canada and the United Kingdom (where child poverty has been of major concern) and India (where similar household arrangements prevail) and the status and trends of child poverty are found to be very different. As for mobility, applying new techniques for measuring its degree we observe that the life chances of children born under its regimen have improved substantially but become increasingly dependent upon their parental circumstances. Thus, consistent with the increased parental investment per child that the OCP engendered, there is a much closer association between the characteristics of subsequent generations or a substantial reduction in generational mobility. This phenomenon is found to be particularly prevalent in the lower income quantiles reinforcing a dynastic notion of poverty.

Introduction.

One of the most controversial and far reaching population control policies in recent history is China's One Child Policy (OCP) introduced in 1979. Directed at China's large population growth rate, the OCP represented a considerable intervention in the household choice process, by fines and various other forms of coercion families were encouraged to limit the production of offspring. The intervention changed fundamentally the nature of both existing and anticipated marriage arrangements and influenced family formation decisions in many dimensions. Anderson and Leo (2007), in studying the impact of the policy on family formation, construed the OCP as a rationing policy constraining the quantity (but not the quality) of children and, following Neary and Roberts (1980) and Deaton (1981), anticipated that the demand equations for quantity and quality of children would be affected accordingly. Increased positive assortative pairing of couples was observed as was increased investment in children and, also consistent with rationing theory, income ceased to be a factor in determining family size, it did however become an increasingly important determining factor in investment in children. Such changes in family formation behavior may well have had substantial impacts on the wellbeing of children both in the context of child poverty and their generational mobility.

Child poverty has been a major issue in western societies where it is strongly associated with single parent family situations (in the United Kingdom children of lone parents are subject to more than twice the risk of poverty compared to children of couples (Brewer et. al. (2006a))). Its eradication has been a policy target in both Canada and the United

Kingdom[1], where child poverty figures of the order of 1 in 6 and 1 in 4 respectively have been cited in recent years (as compared to 1 in 10 and 1 in 7 respectively for people in poverty in the population at large). Though child poverty has not been such a direct policy target in the United States (where roughly speaking 1 in 5 children are in the poor group as compared with 1 in 10 households in the population at large) policies promoting generational mobility, reducing the dependence of child outcomes on parental circumstances, have been pursued. This can be seen as an attack on the dynastic nature of poverty (Kanbur and Stiglitz (1989)) wherein children of the poor have a greater likelihood of being members of the poor club when they become adults. In formalizing the imperative, Roemer (2002) remarked that equal opportunity is "Probably the most universally supported conception of Justice advanced in societies..". With roots in recent egalitarian political philosophy (See Dworkin (1981)) the concept sees differential outcomes as ethically acceptable when they are the consequence of individual choice and action but not ethically acceptable when they are the consequence of circumstances beyond the individual's control.

Here we look at the impact of the OCP on child poverty and generational mobility in the context of the parent/child educational attainments. Stochastic Dominance techniques are employed to compare implicit child and adult income size distributions and new

[1] Without being very explicit in 1989 the Canadian House of Commons unanimously resolved to "seek to achieve the goal of eliminating poverty among Canadian children by the year 2000" (to date a little progress has been made though children are still over-represented in the poverty group). With a similar vagueness in 2000 the British Government vowed to "half child poverty by 2010 and eliminate it by 2020", to this end the UK government committed £8 billion to child contingent income support in the 1999-2004 period (Brewer et. al. (2006)) and some benefit has accrued, by 2004 the number of children in poverty had fallen by some 700,000 and the child poverty rate was at its lowest since the 1980's. These heroic intentions were not matched by the United States administration where the policy focus has been more on improving generational mobility (e.g. "No Child Left Behind").

techniques for measuring the degree of mobility are applied to a data set on urban households from 6 provinces. The impact of the policy on the life chances of children born under its regimen in terms of child poverty and generational mobility is observed using samples of family cohorts for the years 1987, 1992, 1997 and 2001 drawn from an urban survey of six provinces in China, Shaanxi, Jilin, Hubei, Sichuan, Guangdong and Shandong[2]. The first two provinces may be considered low, the second two intermediate and the third pair high income provinces.

Table 1. GDP Comparison

Unit: billion RBM Yuan

	1992	2001
Shanxi	33.96	133.33
Jilin	36.65	149.14
Hubei	69.07	339.89
Sichuan	87.61	354.26
Guangdong	127.18	751.88
Shandong	128.06	678.63

Note: data of 1992 is the average GDP from 1985 to 1992, that of 2001 is from 1994 to 2001

Source: SSB

To anticipate some of the results, unlike the western experience, children do not appear to be over-represented in the poverty group either before or after the OCP, indeed the OCP seems to have strengthened the sense in which they are under-represented. However the OCP does appear to have substantially reduced generational mobility, especially among the lower quantiles of the income distribution. In section 1 the impact of OCP on the nature of child poverty is examined and compared with changes in the child poverty situation in the UK, Canada and India. Section 2 considers how increased investments in

[2] These data were obtained from the National Bureau of Statistics as part of the project on Income Inequality during China's Transition organized by Dwayne Benjamin, Loren Brandt, John Giles and Sangui Wang.

children have impacted their educational attainment and Section 3 considers the impact of the OCP on generational mobility in China. Finally some conclusions are drawn in section 4.

Section 1. Child Poverty and Wellbeing.

In western economies concern with child poverty is concern with an over-representation of children in the poverty "club", alternatively put the children in those societies experience a greater degree of measured poverty than do the adults. How this poverty "club" is measured has been the subject of much debate. What should the poverty frontier be? Should incidence, depth or intensity measures be used? How should adults and children be compared within the context of a household? To what extent are there returns to scale in household consumption? All are issues that have received much attention. To this list should be added the question of the household sharing rule which permits the identification of child and adult income distributions? Thanks to results in Atkinson (1987) many of these debates (aside from the last few) may be circumvented by employing stochastic dominance techniques.

Atkinson's result, also noted in Foster and Shorrocks (1988), establishes a useful link between poverty indices and stochastic dominance relations for more general welfare comparisons which may be summarized as follows. If income distribution $f_A(x)$ stochastically dominates income distribution $f_C(x)$ over the interval 0 to x* at a particular order then all poverty measures in a specific class will record greater poverty for society C than society A for any poverty line up to x*. Intensity of poverty measures require

dominance of order three or less, depth of poverty measures require dominance of two or less and incidence measures (e.g. poverty rates) require dominance of order 1. Noting that dominance of order j implies dominance of order k for k > j we see that first order dominance of A over C implies greater poverty in C than in A for any poverty measure based upon a cut-off < x*. Thus in the present context child $f_C(x)$ and adult $f_A(x)$ income distributions can be compared and if the former is stochastically dominated by the latter at all incomes less than x* we can conclude that children are over represented in the poverty group however it is measured as long as the poverty cutoff is less than x*.

Child and adult income distribution comparisons, computed on the basis of attributing the household equivalized income (using the square root rule alluded to earlier) to each child or adult in the household[3] are made for our Chinese datasets for the years 1987 and 2001. Similar comparisons were calculated for the United Kingdom for 2002 and 1996, Canada for 2004 and 1997 and India for 2004 and 1994[4]. Summary statistics for the comparison distributions are reported in Table 2. We would briefly note that the location measures for the child distributions are greater than the location measures for adult distributions in China whereas the situation is reversed in the UK, Canada and India (note that all comparisons are in within country real terms). Child distributions are always less dispersed than the corresponding adult distributions in all comparisons. With the exception of the 1987 Adult distribution in China, which possesses a much larger

[3] Comparing child and adult income distributions in this way involves very strong implicit assumptions about the way income is allocated or shared within the family which is fundamentally an unobservable phenomenon. Different sharing rules would produce substantially different outcomes establishing what those rules may be is a matter for ongoing research (see Browning et. al. (2006) for instance).

[4] The non China data used in this study relate to United Kingdom household income net of direct taxes and is drawn from the annual Family Resources Survey , for Canada the Statistics Canada Survey of Household spending is used and for India NSS rounds 50 and 60 from the Ministry of Statistics.

standard deviation compared to its Child counterpart, relative variability is pretty stable across distributions.

Table 2. Sample characteristics log adult real equivalized household income.

	Mean	Median	Standard Deviation	Coefficient of variation	Sample Size
China Child 1987	4.7819	4.8548	0.6164	0.1289	3387
China Adult 1987	4.5317	4.8040	1.2556	0.2771	2074
China Child 2001	8.8374	8.9197	0.8642	0.0978	10569
China Adult 2001	8.6875	8.7916	0.9050	0.1042	3936
UK Child 1996	5.3423	5.3390	0.6380	0.1194	17386
UK Adult 1996	5.5021	5.5277	0.6797	0.1235	60323
UK Child 2002	5.5985	5.6089	0.6750	0.1206	9401
UK Adult 2002	5.7257	5.7703	0.7509	0.1311	32691
Canada Child 1997	10.0136	10.0858	0.6685	0.0668	13040
Canada Adult 1997	10.1234	10.1721	0.6952	0.0687	34594
Canada Child 2004	10.3216	10.3658	0.6746	0.0654	9214
Canada Adult 2004	10.3731	10.4193	0.7115	0.0686	26027
India Child 1994	11.3268	11.2815	0.5446	0.0481	67033
India Adult 1994	11.4490	11.4012	0.5748	0.0502	141215
India Child 2004	11.9622	11.9234	0.4811	0.0402	12342
India Adult 2004	12.0444	11.9996	0.5151	0.0429	13716

Table 3 reports the comparisons of the distributions. It is readily seen that for 2001 at every decile cut-off up to the 9th there were smaller proportions of children in the group than there were adults, very much a property of the child's income distribution first order dominating that of the adults. Thus we can safely conclude that for almost any poverty measure at any cutoff line child poverty would be less than adult poverty. The same is almost true for the 1987 year where the proportion of children is always less than the proportion of adults for every decile cut-off up to the 7th, so that child poverty would be less than adult poverty for any poverty line up to the 7th decile for virtually all poverty measures[5]. Tables 3a, 3b and 3c report the corresponding comparisons for selected years

[5] Actually 1st order dominance would be rejected over the whole income range since for the 8th and 9th deciles child shares are significantly greater than adult shares but second order dominance of the adult income distribution by the child's distribution would prevail (see appendix) so that all depth and intensity poverty measures would record less poverty for children than for adults over the whole income range.

for Canada, the United Kingdom and India. Notice that for the U.K. and India comparators the reverse is true, at every decile cut-off up to the 9th the child's income

Table 3. Child and Adult proportions at income deciles: Urban China.

Case	Decile	Cut-off	Child Share	Adult Share	Difference	Diff Std Err
China	1	0.8603	0.0880	0.1319	-0.0439	0.0105
2001	2	0.9088	0.1825	0.2477	-0.0652	0.0136
N_c=10569	3	0.9506	0.2799	0.3537	-0.0738	0.0152
N_a=3936	4	0.9848	0.3792	0.4558	-0.0766	0.0160
	5	1.0113	0.4837	0.5434	-0.0598	0.0161
	6	1.0339	0.5913	0.6235	-0.0322	0.0157
	7	1.0555	0.6932	0.7180	-0.0248	0.0147
	8	1.0831	0.7911	0.8239	-0.0328	0.0125
	9	1.1236	0.8916	0.9223	-0.0307	0.0091
China	1	0.8860	0.0765	0.1389	-0.0624	0.0089
1987	2	0.9402	0.1760	0.2406	-0.0646	0.0114
N_c=3387	3	0.9824	0.2772	0.3375	-0.0603	0.0129
N_a=2074	4	1.0131	0.3770	0.4378	-0.0608	0.0137
	5	1.0394	0.4804	0.5352	-0.0548	0.0139
	6	1.0645	0.5890	0.6176	-0.0286	0.0136
	7	1.0906	0.6994	0.7049	-0.0056	0.0127
	8	1.1192	0.8087	0.7869	0.0218	0.0112
	9	1.1633	0.9114	0.8809	0.0305	0.0086

Table 3a. Child and Adult proportions at income deciles: Canada.

Case	Decile	Child Share	Adult Share	Difference	Diff Std Err
Canada	1	0.1107	0.0966	0.0141	0.0037
2004	2	0.2018	0.2051	-0.0034	0.0049
N_c=10569	3	0.3062	0.2994	0.0068	0.0056
N_a=3936	4	0.4145	0.3954	0.0190	0.0060
	5	0.5247	0.4926	0.0321	0.0061
	6	0.6322	0.5903	0.0419	0.0059
	7	0.7418	0.6862	0.0556	0.0054
	8	0.8523	0.8000	0.0523	0.0044
	9	0.9239	0.8927	0.0312	0.0034
Canada	1	0.1283	0.0903	0.0380	0.0033
1997	2	0.2274	0.1936	0.0338	0.0042
N_c=3387	3	0.3301	0.2891	0.0410	0.0048
N_a=2074	4	0.4428	0.3847	0.0581	0.0051
	5	0.5469	0.4838	0.0631	0.0051
	6	0.6545	0.5807	0.0738	0.0049
	7	0.7558	0.6794	0.0764	0.0045
	8	0.8510	0.7817	0.0693	0.0038
	9	0.9360	0.8919	0.0441	0.0027

Table 3b. Child and Adult proportions at income deciles: United Kingdom.

Case	Decile	Child Share	Adult Share	Difference	Diff Std Err
2002	1	0.1191	0.0946	0.0245	0.0037
N_c=9401	2	0.2564	0.1839	0.0725	0.0050
N_a=32691	3	0.3855	0.2754	0.1101	0.0056
	4	0.4889	0.3745	0.1144	0.0058
	5	0.5975	0.4721	0.1254	0.0058
	6	0.6906	0.5741	0.1165	0.0055
	7	0.7816	0.6766	0.1050	0.0050
	8	0.8641	0.7816	0.0825	0.0042
	9	0.9349	0.8900	0.0449	0.0031
1996	1	0.1313	0.0910	0.0404	0.0028
N_c=17386	2	0.2805	0.1768	0.1036	0.0037
N_a=60323	3	0.3977	0.2718	0.1259	0.0041
	4	0.4987	0.3716	0.1271	0.0043
	5	0.6041	0.4700	0.1341	0.0042
	6	0.7055	0.5697	0.1358	0.0040
	7	0.7924	0.6734	0.1191	0.0036
	8	0.8697	0.7800	0.0897	0.0030
	9	0.9370	0.8894	0.0476	0.0022

Table 3c. Child and Adult proportions at income deciles: India.

Case	Decile	Child Share	Adult Share	Difference	Diff Std Err
2004	1	0.1095	0.0915	0.0180	0.0037
	2	0.2173	0.1845	0.0328	0.0050
	3	0.3263	0.2766	0.0497	0.0057
	4	0.4300	0.3732	0.0568	0.0061
	5	0.5343	0.4697	0.0646	0.0062
	6	0.6374	0.5663	0.0711	0.0061
	7	0.7375	0.6663	0.0712	0.0056
	8	0.8294	0.7735	0.0559	0.0049
	9	0.9187	0.8832	0.0354	0.0037
1994	1	0.1210	0.0901	0.0309	0.0015
	2	0.2383	0.1818	0.0564	0.0019
	3	0.3517	0.2755	0.0762	0.0022
	4	0.4599	0.3716	0.0883	0.0023
	5	0.5629	0.4701	0.0928	0.0023
	6	0.6626	0.5703	0.0923	0.0023
	7	0.7563	0.6733	0.0830	0.0021
	8	0.8448	0.7787	0.0661	0.0018
	9	0.9277	0.8869	0.0408	0.0013

share is greater than the adult's (both in 1996 and 2002), very much a characteristic of the adult income distribution stochastically dominating the child's income distribution at all

orders indicating an overrepresentation of children in the poverty group however defined. The same is true for Canada except for the 2nd decile in the 2004 comparison (but this difference is not significantly different from 0 and not enough to contradict the same dominance result as for the U.K. and India)

In essence the comparisons which permitted such strong and sweeping statements to be made were of the form:

$$\int_0^x (F_C^i(z) - F_A^i(z))dz \leq \quad 0 \quad \forall \quad x$$

$$where:$$

$$F_H^{i+1}(x) = \int_0^x F_H^i(z)dz \; with \; F_H^0(z) = f_H(z) \; for \; H = A, B \; and \; i = 1, 2, ..$$

with the strict inequality holding at least somewhere. These are conditions under which we can infer that society C is better off than society A. Thus from the above, following Foster and Shorrocks (1988), we may infer that the society of children were better of than the society of adults for all utilitarian social welfare functions in 2001 and for all social welfare functions that express a preference for mean preserving progressive transfers in 1987. The reverse is true for the UK, Canada and India in both their comparison periods. The society of adults is better off than the society of children in terms of all utilitarian social welfare functions. However we can take the analysis further, consider the income distribution of society H in year k to be $f_{Hk}(x)$ and consider the condition:

$$\int_0^x [(F_{C2001}^i(z) - F_{A2001}^i(z)) - (F_{C1987}^i(z) - F_{A1987}^i(z))]dz \leq \quad 0 \quad \forall \quad x$$

This responds to the question "Does the extent to which the child's society was better off than the adults in 2001 dominate the extent to which the child's society was better off than the adult's in1987?". This is essentially a difference in dominance comparison, part of the toolkit for studying polarization (Anderson (2004)). The comparison results employing the Wolak (1989) method for comparing multivariate inequalities are reported in Table 4 for China 2001 and 1987, for Canada 2004 and 1997 for the UK 2002 and 1996 and for India 2004 and 1994. In the case of all of the countries the notion that the change in the child's distribution dominated that of the adult's distribution is not rejected whereas the notion that the change in the adult's distribution dominated that of the children is strongly rejected. Because the initial conditions differ, this has different implications for China as compared to the other countries. For China it means that the child's income distribution has moved further away from the Adults income distribution, child and adult societies have become more polarized as it were, whereas for Canada, India and the United Kingdom it means that child and adult groups in those societies have moved closer together or depolarized, children are catching up here in a welfare sense.

Table 4 Difference in differences comparisons, upper tail probabilities.

Comparison	Adult Change Dominates Child Change	Child Change Dominates Adult Change
China 1987-2001	0.0000	0.9409
Canada 1997-2004	0.0000	0.8925
United Kingdom 1996-2002	0.0000	0.9007
India 1994-2004	0.0000	0.9305

It is thus safe to conclude that, on the basis of our samples, relative to adult poverty, child poverty is not the issue in China that it appears to be in the other societies under comparison, indeed one may ask why there is no concern over increasing relative adult poverty!

Section 2. The OCP and Investment in Children

One feature of the OCP impact on family formation noted in Leo and Anderson (2007) is the increased investment in child quality. To illustrate the issue here investment in children as reflected in their educational attainments is considered. For comparison purposes only children over 20 are studied so that the parental aspect of the investment activity may assume to have been completed and cohorts of such children for the years 1987 and 2001 in the six provinces are investigated. The salient features of the data together with a comparison of the educational attainment cdf's are reported in table 5.

Table 5. Educational Attainment and household characteristics (children over age 20) raw data

Means	1987	2001
Age	24.7346	26.8895
%/100 female	0.4904	0.4650
Educational Status*	2.8524	3.7316
Birth Order	1.3854	1.2624
Family Size	4.4989	3.8456
Family Income (nominal)	323.1836	24616.638
Fathers Educational Status	2.5605	3.4118
Mothers educational Status	1.6582	2.4557
Fathers Age	49.2081	51.5134
Mothers Age	50.1348	53.0869
Number of Siblings	2.3567	1.6068
Number of Observations.	942	1185

First order dominance comparison.

	Primary or less	Middle School	High school	Technical School	College and higher
1987 cdf	0.03175	0.32804	0.85608	0.91640	1.00000
2001 cdf	0.01266	0.19072	0.50042	0.66751	1.00000
Difference	0.01909	0.13733	0.35566	0.24889	
(Std Error)	(0.00656)	(0.01907)	(0.01848)	(0.01638)	

*The Education level variable for 1987 is not strictly comparable with 2001 though it has been adjusted for the dominance comparisons

Note the dramatic improvement in children's educational status across the two cohorts with the 2001 cohort attainment distribution strongly first order dominating that of 1987.

To examine the determinants of child quality educational attainment is regressed upon the logarithm of adult equivalized household income, father's and mother's educational status, household size, and the gender of the child (1 if female 0 otherwise) and birth order (representing the order in which the child arrived in the family, 1 = first child, 2 = second child…). Household income equivalization is based upon the square root rule (Brady and Barber (1948)) under which household income is deflated by the square root of the number of persons in the household and reflects returns to scale in household consumption. Parental educational status is included to reflect both inherited abilities and parental preferences, household size is included independently of income to reflect investment scale effects beyond the consumption nexus. Gender and birth order effects are included following Bjorklund et. al (2004) and Kantarevic and Mechoulan (2006).

Table 6 reports regressions for 1987 and 2001 cohorts which were performed for single child, two child and more than two child family situations noting that they are really mixtures of pre and post OCP families where the post OCP family mixing coefficient is increasing with the cohort year so that the proportion of pre OCP children in the samples upon which each regression is based will diminish as time progresses. The regression equations for the two cohorts are very different. Broadly speaking the impact of household income has increased significantly over time as has the combined effect of mother's and fathers educational status (a combined coefficient of 0.24 in 1987 moving to 0.26 in 2001 for single children, 0.16 to 0.19 for the two child family comparison and 0.16 to 0.58 for the more than two child comparison) and the mothers status has become more important relative to fathers status. A significant negative birth order effect emerged in two child families together with a positive birth order effect more than two

child families. Finally the positive gender effect - favoring girls - that emerged (significant in 2001) should be noted. The overall picture is one of increasing dependence on parental characteristics in terms of income and educational status.

Table 6 Child Quality Regressions

2001 (n= 1185)

	One Child Families		Two Child Families		> Two Child Families	
	Coefficient	Std Error	Coefficient	Std Error	Coefficient	Std Error
Constant	0.5096	(0.4828)	-0.6600	(0.5787)	4.0144	(0.7619)
Income	0.4845	(0.0603)	0.6824	(0.0641)	-0.2779	(0.0903)
Father edu	0.0885	(0.0218)	0.0556	(0.0195)	0.2165	(0.0244)
Mother edu	0.1743	(0.0254)	0.1377	(0.0226)	0.3634	(0.0313)
Family size	0.0486	(0.0611)	-0.0433	(0.0547)	0.0481	(0.0487)
gender	0.1535	(0.0593)	0.1474	(0.0675)	0.0281	(0.0570)
birth order			-0.1681	(0.0682)	0.0946	(0.0335)
provincial	-0.1684	(0.1125)	0.1844	(0.1003)	1.0796	(0.1742)
dummies	-0.4085	(0.1466)	-0.7306	(0.1436)	0.4493	(0.2238)
	-0.3821	(0.1386)	-0.3536	(0.1349)	0.9987	(0.2083)
	-0.6808	(0.1817)	-0.8753	(0.1756)	1.6418	(0.2572)
	-0.6685	(0.1288)	-0.8824	(0.1295)	1.0643	(0.1779)
R^2 StdE	0.6332	0.4656	0.6099	0.5454	0.9620	0.0518

1987 (n= 945)

	One Child Families		Two Child Families		> Two Child Families	
	Coefficient	Std Error	Coefficient	Std Error	Coefficient	Std Error
Constant	2.6593	(0.2281)	2.5403	(0.3181)	4.1953	(0.2507)
Income	0.3008	(0.0313)	0.4790	(0.0514)	0.0499	(0.0475)
Father edu	0.1389	(0.0179)	0.0391	(0.0173)	0.0320	(0.0090)
Mother edu	0.0955	(0.0188)	0.1196	(0.0195)	0.1253	(0.0211)
Family size	0.1460	(0.0556)	0.0377	(0.0556)	-0.0025	(0.0363)
gender	0.0928	(0.0606)	-0.0887	(0.0616)	-0.0029	(0.0509)
birth order			-0.0036	(0.0675)	-0.0252	(0.0305)
provincial	-0.0391	(0.1218)	0.2817	(0.1136)	0.1910	(0.0882)
dummies	-0.6188	(0.0926)	-0.8605	(0.1306)	-0.0372	(0.1054)
	-0.5412	(0.1014)	-0.5022	(0.1268)	0.0966	(0.1021)
	-0.3481	(0.1066)	-0.5955	(0.1229)	-0.0976	(0.0992)
	-0.6436	(0.1052)	-0.7743	(0.1446)	-0.3006	(0.1126)
R^2 StdE	0.8704	0.1316	0.6367	0.3690	0.6441	0.2251

3. The One Child Policy and Generational Mobility.

Greater parental investments in child quality undoubtedly strengthen the ties between generational income distributions, making it is more likely that parents with high (low) incomes will have children who will earn high (low) incomes when they become adults. This increased dependence between child and parent outcomes constitutes a reduction generational mobility. Here the generational mobility phenomenon is examined by studying the dependence between parent incomes and educational attainments and child educational attainments.

Mobility has mostly been studied in the context of intergenerational income transitions (see Corak (2004) (2006) for a survey and examples) generally assessed via the regression coefficient (β) of the child income (y) on the corresponding parent income (x) inferring mobility (equal opportunity) as $\beta \rightarrow 0$ and immobility (unequal opportunity) as $\beta \rightarrow 1$. However mobility interpretations of β are limited by its connection to the correlation coefficient ρ_{yx} ($\beta = \rho_{yx}(\sigma_y / \sigma_x)$), and that statistics ability to reflect dependence, they are further circumscribed by the degree to which the y and x relationship is homogeneously linear across all income strata[6]. However if the degree of dependence is the issue and the setting is not homogeneous, the transition matrix between the common quantiles of the marginal distributions f(x) and f(y) can be more informative. This has given rise to the application of techniques derived from Markov Chain processes

[6] As an index β would not prove very effective if immobility were just confined to the lowest income group for example. Indeed there are dangers with interpreting zero correlation with perfect mobility, imagine a deterministic world (perfectly immobile) where below a certain parental income there is an exact negative relationship between parent and child outcomes whereas above that income there is an exact positive relationship between parent and child outcomes, an appropriately balanced sample would yield 0 correlation with an inferred perfect mobility for what is a completely deterministic and immobile state.

and the development of mobility indices, some based upon the nature of the transition matrix directly (with complete mobility the columns of a transition matrix are identical), some based upon other related concepts[7]. When the quantiles or categories of f(x) and f(y) match or are common, such an analysis is straightforward, but when they are not, it is not. Here, since the data of necessity does not present in terms of the common quantiles of two single variables, more general transition processes need to be contemplated between generations characterized by different sets of characteristics which do not readily lend themselves to a Markov chain interpretation.

Following Anderson and Leo (2006) the extent to which independence accords with the data can be indexed by an overlap measure given by:

$$OV = \sum_i \sum_j \min(p_{ij}^o, p_{ij}^e)$$

Where p_{ij}^o corresponds to the observed ij'th cell probability and p_{ij}^e corresponds to the expected ij'th cell probability under the null hypothesis of independence where i corresponds to the i'th child characteristic configuration and j corresponds to the j'th parental characteristic configuration. This measure forms a very natural index since it reflects the proximity of the data to the hypothesis of interest. When the data completely conform to the hypothesis of interest (in this case perfect mobility) OV = 1, otherwise $0 \leq OV < 1$. OV is easily calculated, p_{ij}^o is simply estimated from the observed cell sample proportions and p_{ij}^e is estimated from the product of the corresponding empirical marginal proportions. An attractive feature of these indices is that they can be readily

[7] Bartholemew (1982), Blanden et. al. (2004), Chakravaty (1995), Dearden et. al. (1997), Hart (1983), Maasoumi (1986), Maasoumi and Zandvakili (1986), Prais (1955), Shorrocks (1976), (1978) have all produced "Transition" based mobility indices many of which are discussed in Maasoumi (1995)).

applied when the underlying parent child transition matrices are not square and transitions are between multivariate environments. For example given continuously measured parental characteristics w and x with joint density f(w,x) and continuously measured child characteristics y and z with joint density g(y,z) and a joint density of all characteristics given by h(w,x,y,z) then OV estimates the magnitude ∫∫∫∫min[h(w,x,y,z),(f(w,x)g(y,z))]dwdxdydz. In addition they appear to have asymptotically normal sampling distributions[8], conveniently facilitating inferences about trends toward or away from independence over time. These indices can be more focused, concentrating on a subset of cells that relate to particular features of interest. So for example mobility amongst the poor could be examined by specifying a null in which only independence with respect to the poor is entertained so that the mobility of the i'th subgroup can be considered in terms of:

$$OV_i = \sum_{j=1}^{k} \min((\frac{p_{ij}}{p_{i.}}), p_{.j})$$

Where $p_{i.}$ and $p_{.j}$ are marginal row and column probabilities respectively[9].

To study the impact on generational mobility of the OCP changes in the degree of dependence between parent and child characteristics pre and post OCP are considered. To identify pre and post OCP children the age of the mother at the introduction of the one child policy is taken into consideration. Households where the mother was less than 25 at the introduction of OCP are considered to have made their parenting decisions in the

[8] The distribution of OV_{Ind} can be shown to be asymptotically normal by noting that, under the null of independence, both p_{ij}^{o} and p_{ij}^{e} are normal with means p_{ij} and variances $p_{ij}(1-p_{ij})/n$ and, following results in Daganzo (1980) based on Clark (1961), $\min(p_{ij}^{o}, p_{ij}^{e})$ will also be normal, and OV_{Ind}, being a sum of such terms, will also be asymptotically normal. Anderson, Ge and Leo (2006) presents a small Monte Carlo exercise supporting normality of overlap indices.

[9] This possibility calls for a concept of "Qualified Equal Opportunity" or "Conditional Mobility" which has been developed elsewhere (Anderson, Leo and Muelhaupt (2007)).

context of being an OCP family. Households where the mother was over 35 at the onset of the OCP are considered to have made their parenting decisions prior to the OCP. The parent's characteristics are household income and educational attainment (the maximum

Table 7. Mobility Differences Pre and Post OCP families

Children 20-30.	Mother's Age at OCP		
	<25 at OCP (Post OCPDeciders) Index (Std Err.)	>35 at OCP (Pre OCP Deciders) Index (Std Err.)	Difference [P-Value]
Mobility all Children, Education-Education n=236/1356	0.78244 (0.02686)	0.84834 (0.00974)	-0.065895 [0.02108]
Mobility for all Children, Education-Income	0.88279 (0.02094)	0.93245 (0.00682)	-0.049655 [0.024135]
Mobility for all Children, Education-Education-Income	0.69844 (0.02987)	0.81828 (0.01047)	-0.11984 [0.000153]
Mobility of Males, Education-Education* n=114/746	0.80371 (0.03720)	0.85097 (0.01304)	-0.047264 [0.23052]
Mobility of Females, Education-Education n=122/610	0.73206 (0.04010)	0.83475 (0.01504)	-0.10260 [0.01649]
Mobility of Males, Education-Income	0.85488 (0.03299)	0.93218 (0.00921)	-0.077303 [0.024003]
Mobility of Females, Education-Income	0.85038 (0.03229)	0.92247 (0.01083)	-0.072091 [0.034302]
Mobility of Males, Education-Education-Income	0.6385 (0.04500)	0.81126 (0.01433)	-0.17276 [0.000254]
Mobility of Females, Education-Education-Income	0.64271 (0.04339)	0.7945 (0.01636)	-0.15179 [0.001062]
Mobility 1st born, Education-Education N=187/663	0.77912 (0.03034)	0.85884 (0.01352)	-0.079723 [0.016381]
Mobility 2nd born, Education-Education* 42/481	0.75624 (0.06625)	0.84192 (0.01663)	-0.085682 [0.20971]
Mobility 1st born, Education-Income	0.85982 (0.02539)	0.93219 (0.00976)	-0.072374 [0.007797]
Mobility 2nd born, Education-Income*	0.88209 (0.04976)	0.90989 (0.01306)	-0.027799 [0.58896]
Mobility of 1st born, Education-Education-Income	0.69112 (0.03379)	0.80771 (0.01531)	-0.11659 [0.001670]
Mobility of 2nd born, Education-Education-Income	0.49302 (0.07714)	0.7778 (0.01896)	-0.28478 [0.000337]

of the parent's attainments) the child's characteristics are their educational attainments[10]. Only children between the ages of 20 and 30 were considered so as to allow for them to have completed their education. In addition to overall child - parent mobility, mobility by gender and birth order of child are considered following Bjorkland et.al. (2004). Both univariate (child achievement – parent achievement and child achievement – parent income) multivariate indices (child achievement – parent achievement and income) are employed. As table 7 reveals with very few exceptions mobility diminished significantly in virtually every situation that was examined, the three exceptions were the parent's education – child's education comparison for males and the univariate comparisons for the 2nd born[11].

Table 8. Educational mobility by income quartile.

	Mother's Age at OCP		
	<25 at OCP (Post OCPDeciders)	>35 at OCP (Pre OCP Deciders)	Diff/P-Value
First Quartile	0.70321 (0.06810) 45	0.86021 (0.01376) 635	-0.1570 [0.023843]
Second Quartile	0.73291 (0.05531) 64	0.85902 (0.02003) 302	-0.1261 [0.032031]
Third Quartile	0.81564 (0.05136) 57	0.84537 (0.02138) 286	-0.0297 [0.59301]
Fourth Quartile	0.72204 (0.05355) 70	0.77398 (0.03627) 133	-0.051943 [0.42187]

[10] Educational attainment is measured as an integer indexed from 0 to 5 with 5 = college graduates and above, 4 = technical secondary school, 3 = high school, 2 = middle school, 1 = primary school and lower.

[11] The education – education results by gender bear comparison with those in Canada over a similar time frame (Anderson, Leo and Muelhaupt (2007)) There male mobility also remained constant whereas female mobility rose in part due the greater investments in girls by those poor in circumstance, a result of the move toward gender equity in the last part of the last century which in terms of children in China does not appear to have been an issue (see appendix 2).

A closer look at education-education mobility by income quartile is reported in table 8 and the results are somewhat more troublesome with the only significant reductions in the two lowest income quartiles. Thus conditional mobility appears to have diminished predominantly amongst the poor corresponding to diminished life chances for the children of poor families from escaping that predicament.

Section 4. Conclusions.

Unlike some countries in the west, children do not appear to be over-represented in the poverty group in Urban China neither before nor after the one child policy. What may be interpreted as the child's income distribution is seen to stochastically dominate the adult income distribution in the pre One Child Policy environment at the second order. The post one child policy result is even stronger with the child income distribution first order dominating that of the adult distribution. Indeed it appears that the OCP has polarized (i.e. widened the gap between) child and adult income distributions[12]. Thus the general dominance relationship between child and adult income distributions does not appear to be a result of the OCP[13] though not surprisingly it does appear to have precipitated an improvement in the wellbeing of children relative to the adult population.

With respect to mobility, the increased intensity of investment in child quality brought about by the one child policy has reinforced the link between parent and child quality and

[12] On the other hand policies pursued in Canada, the UK and India, our comparator societies, appeared to have narrowed the gap between the child and adult income distributions where generally the adult income distribution dominates that of the children.

[13] It could well be a consequence of the nature of household formation and the extended family found in China (however it does not seem to be the case in Urban India where similar extended family arrangements prevail) together with higher income elasticities of demand for children than are prevalent in the west, all of which is the subject of ongoing research.

reduced generational mobility as a result. This is contrary to the results found for the US for example where generational mobility has increased over the last part of the 20th Century (Anderson and Leo (2006a)). When viewed by income quartile, increases in immobility are found to be more prevalent in the lower income quartiles reinforcing notions of "Dynastic Poverty" discussed in Kanbur and Stiglitz (1986).

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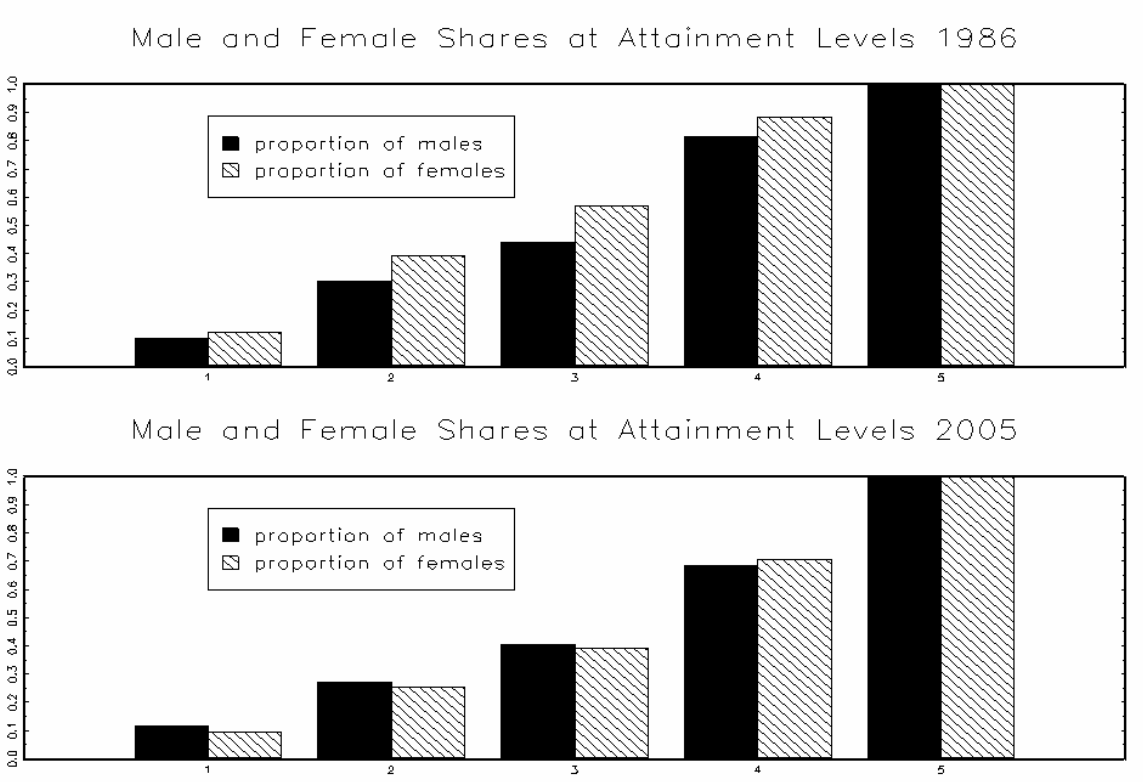
Appendix 1.

Second order dominance results for China 1987.

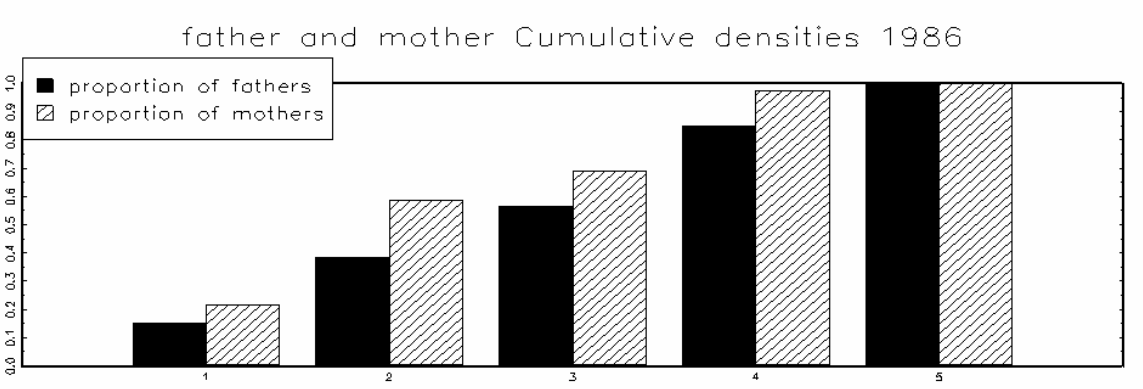
Table 4. Child and Adult proportions at income deciles.

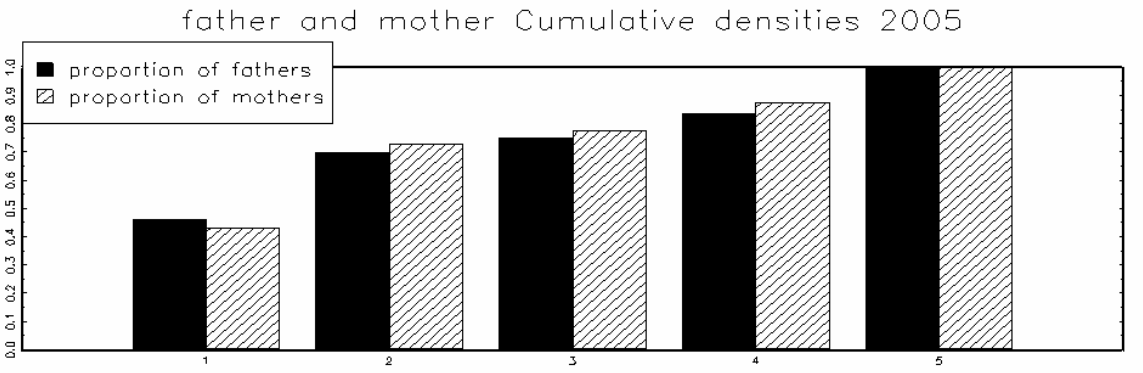
Case	Decile	Cut-off	Child	Adult	Difference	Diff Std Err
China	1	0.8860	0.0153	0.0734	-0.0581	0.0064
1987	2	0.9402	0.0219	0.0837	-0.0617	0.0067
	3	0.9824	0.0314	0.0957	-0.0643	0.0069
	4	1.0131	0.0413	0.1074	-0.0667	0.0071
	5	1.0394	0.0524	0.1201	-0.0677	0.0072
	6	1.0645	0.0658	0.1345	-0.0687	0.0073
	7	1.0906	0.0826	0.1517	-0.0691	0.0075
	8	1.1192	0.1043	0.1732	-0.0689	0.0076
	9	1.1633	0.1425	0.2101	-0.0676	0.0077
	10	1.3860	0.3609	0.4269	-0.0660	0.0078

Appendix 2.Educational Attainment by Gender: A Canadian Comparison

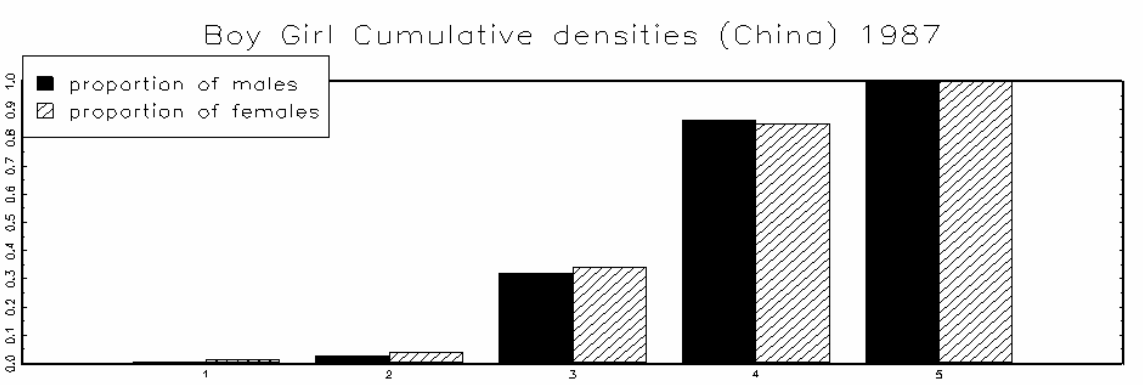


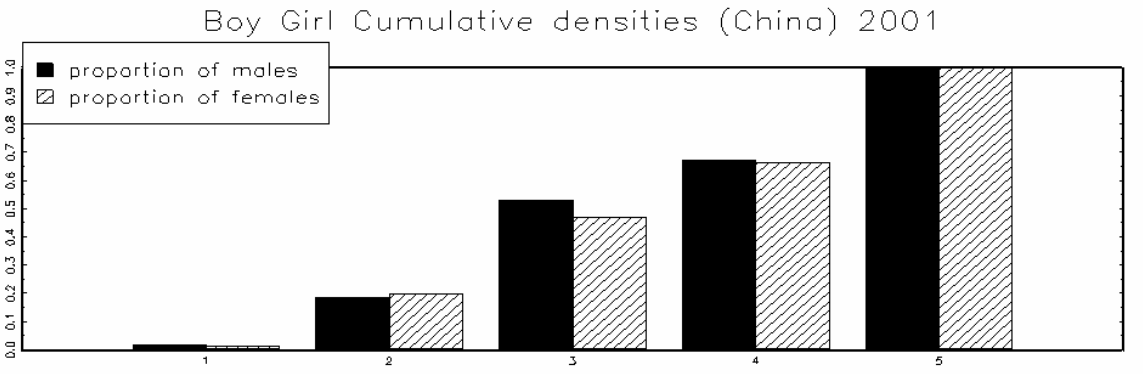
Here it may readily be shown that, whereas the male attainment distribution first order dominates that of females in 1986, by 2005 the female attainment distribution second order dominates that of males. There is a sense in which the gap has more than narrowed women have in fact overtaken men. It is illuminating to compare the attainment levels of the mothers and fathers in the two observation periods by pooling the parents of males and females. What will be observed is a considerable narrowing of the gap between fathers and mothers income distributions mirroring the closing of the gender gap between the attainments of their offspring. All in all it appears that this is a trend that has been continuing for some time.



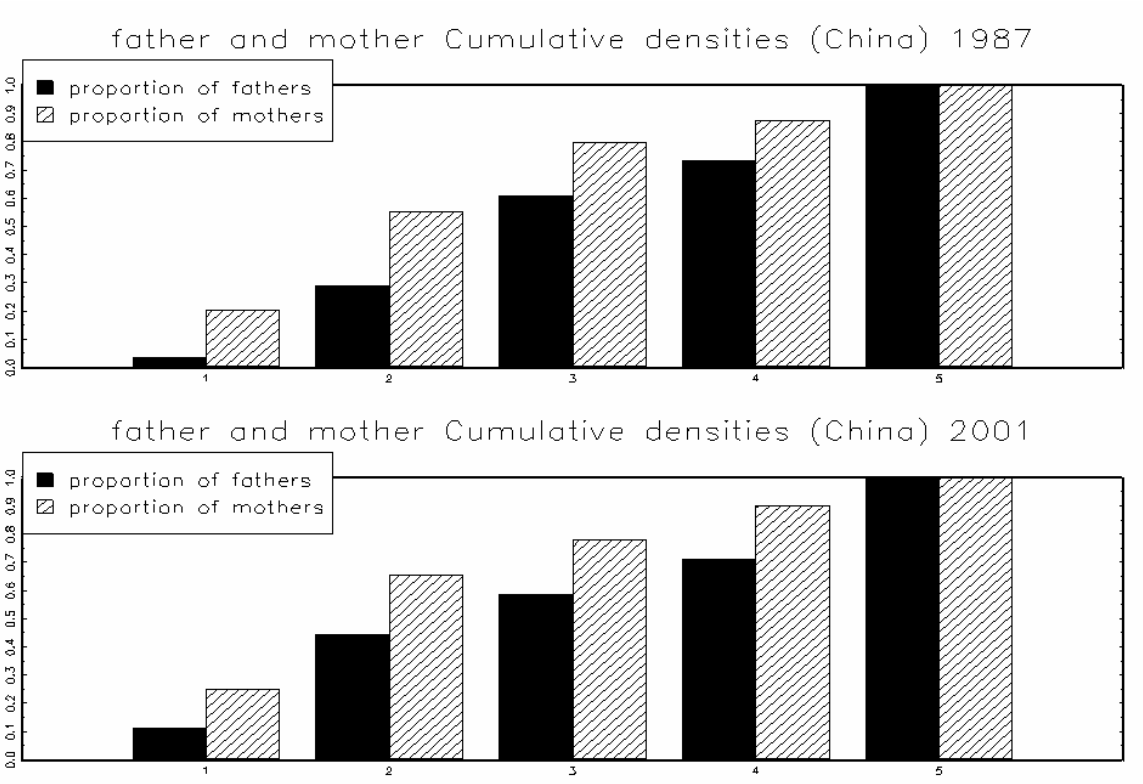


The corresponding diagrams for China are as follows:





It should be noted that the educational levels in this comparison do not correspond to those in the Canadian comparison, nonetheless it remains the case that neither gender dominates the other either at the beginning nor at the end of the period. The same cannot be said for their parents as the following diagram attests:



Here fathers strongly stochastically dominate mothers in their educational attainments. However some caution should be exercised in interpreting these results unlike the children, their parents are not random samples and the strong dominance relationship may be a consequence of sorting behavior. None the less it does give a flavor of what prevailed in earlier generations.