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ECONOMIC RESTRUCTURING AND INCOME INEQUALITY IN URBAN CHINA

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Economic transition from a planned to a market oriented economy is often associated with a widening of income inequality. The nature of this change, however, may differ during different stages of the economic transition. This paper investigates the increase in income inequality in urban China during two phases of economic reform: a moderate reform era (1988–95) and a radical reform era (1995–99). It is found that although income inequality increased considerably during both stages, the nature and causes of the increase are different. In the moderate reform period, the increase in inequality was a result of some parts of society sharing more of the economic gain than others, and the main cause of this inequality is regional income dispersion. During the radical reform period income reductions at the lower end of the distribution is observed, and it is mainly due to the large-scale unemployment generated by labor reallocation.

1. INTRODUCTION

Economic transition from a planned to a market oriented economy is often associated with a widening of income inequality (see Gustafsson and Li, 1997, 1998, 2001; Brainerd, 1998; Milanovic, 1997; Khan *et al.*, 1999; Yang, 1999; Flemming and Micklewright, 2000; among others). Such an increase, however, may be induced by different forces and may affect social stability in different ways. One type of increase in income inequality can be as a result of a reform-induced economic gain that is distributed more to high income than to low income groups. Another type can be as a result of economic restructuring. As a labor market liberalizes, the disparity in the rate of returns to low and high skills may enlarge. In addition, economic restructuring will re-allocate labor from a previously distorted state sector to the market sector, inducing both large-scale unemployment and a sectoral shift in employment (Flemming and Micklewright, 2000). Such a restructuring may reduce incomes at the lower end while increasing incomes at the higher end of the income distribution. In terms of generating social instability, the type of increase in income inequality that reduces incomes at the lower end of the distribution may be particularly worrisome.

In contrast to most Eastern European countries, economic restructuring in China took place gradually. Although it began in the late 1970s, little had changed in urban areas in terms of wage determination, job security, and sectoral

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composition of employment until the late 1980s to the early 1990s (Meng, 2000). This gradual approach may have affected the way in which the increase in income inequality in China differs from that in the Eastern European economies during the initial stages of economic reform. For example, according to Flemming and Micklewright (2000), the significant increase in income inequality in most Eastern European countries was accompanied by income increases at the top together with decreases at the bottom of the distribution. Whereas Zhao and Li (1999) indicate that during the period 1988 to 1995 income for the top 3 percent of households in urban China increased by 53 percent and for the bottom 20 percent increased by 20 percent. Thus, until the mid-1990s, although the income distribution widened everybody was better off. As a result the widening income distribution during this period generated limited social and political concerns.

Since the mid-1990s, however, economic restructuring in urban China has accelerated in the areas of state enterprise reform, social security reform, and labor market reform. As a result, the state and collective employment share has reduced from 76 percent of total urban employment in 1995 to 49 percent in 1999 and unemployment has increased significantly. Although official unemployment figures have been kept very low, at around 3 percent, several different estimates suggest that as many as 15–27 million state sector workers were laid off in 1999, which accounts for an additional 7–12 percent of the urban labor force (Fan, 2000; Appleton *et al*., 2001).

Questions naturally arise as to whether the acceleration of economic restructuring has changed the nature of the increase in income inequality in urban China and to what extent the increase in inequality after the mid-1990s is due to economic restructuring. Using three comparable urban household surveys this paper investigates these questions by comparing the change in income inequality between the periods 1988–95 and 1995–99 and identifying the different contributing factors to the changes in each period. Answers to these questions may add to our knowledge about the relationship between the process of economic transition and income distribution and help policy makers to formulate more appropriate policies to establish social stability.

Previous studies on changes in income inequality in China focused mainly on the period up to the mid-1990s (Knight and Song, 1991; Kahn *et al*., 1992; Aaberge and Li, 1997; Gustafsson and Li, 1997, 1998, 2001; Knight and Li, 1999; Yang, 1999; Khan and Riskin, 2001; Riskin *et al*., 2001). Due to the lack of available data, there are, as yet, no published studies on recent developments. In early 2000, however, the Institute of Economics at the Chinese Academy of Social Sciences, with assistance from the China Statistical Bureau, conducted a new household income distribution survey, which collected information on household income and expenditure in 1999 in six provinces. This paper utilizes this new survey together with two other comparable surveys conducted by the same Institute for the years 1988 and 1995.

In addition to studying the most recent developments in income inequality, this study adopts a new methodology. Most previous studies of income inequality in China adopt methodologies which consider only a limited number of contributing factors (see, for example, Aaberge and Li, 1997; Gustafsson and Li, 1998, 2001). This study employs a regression based decomposition approach developed by Fields (1998) that permits an assessment of a wider range of contributing factors.

The paper is structured as follows. The next section provides background information on the process of economic reform in urban China, its relationship to income distribution, and data and summary statistics on income inequality. Section 3 describes methodology. Section 4 investigates the contributing factors to the level and the change in income inequality. Concluding remarks and policy implications are given in Section 5.

2. Background and Data

China has experienced rapid economic growth since economic reforms began in the late 1970s. During the period 1978–99, per capita real GDP increased by 8.3 percent per year. Household income also increased considerably but varied from period to period. The annual increase in urban household real income was 5.6, 8.3, and 7.3 percent for the periods 1982–88, 1988–95, and 1995–99, respectively (China State Statistical Bureau (SSB), 2000).

Income inequality also increased. The World Bank (1997) reports that the Gini coefficient for China as a whole increased from 28.2 in 1981 to 38.8 in 1995. This increase also varied across different periods of economic reform. Before the early 1990s, economic reform was mainly concentrated on product markets and little was changed in terms of the compressed wage structure, immobility of labor, and the domination of state sector employment (Meng, 2000). Consequently, income distribution changed very slightly and was remarkably egalitarian during this period (Khan and Riskin, 2000).

Factor market reform proceeded gradually from the early 1990s. By the mid-1990s, labor mobility across urban regions and between rural and urban areas increased (see, for example, Meng, 2000; West and Zhao, 2000) and the rate of return to different levels of labor market skills widened (Knight and Song, 2001). Accompanying this reform process, income inequality increased sharply (Khan *et al.*, 1992; Aaberge and Li, 1997; Gustafsson and Li, 1997, 1998, 2001; Knight and Li, 1999; Yang, 1999; Khan and Riskin, 2001; Riskin *et al.*, 2001). Gustafsson and Li (2001) report that the Gini coefficients of household income for urban areas in ten provinces increased from 22.8 in 1988 to 27.6 in 1995. Using the same data Khan and Riskin (2001) report a larger increase in the Gini coefficient from 23.3 to 33.2 for the same period.[1] Increase in regional dispersion is identified as one of the main contributors to the increase in income inequality in this period (see, for example, Gustafsson and Li, 2001; Khan and Riskin, 2001).

Since 1995, urban economic reform has taken a sharp turn. Due to soft budget constraints and other property rights related problems, the Chinese state sector has been performing badly. In 1995–96, around 50 percent of enterprises were making losses. To vitalize the Chinese economy the policy of radical reform for the state enterprises was introduced in 1997 (Appleton *et al.*, 2001). Many small

[1]The reason for this difference may be due to the different definition of income used in these studies. However, by reading the two papers it is not very clear exactly where the difference comes from.

and medium sized loss making state enterprises were bankrupted as a result of this policy and the survivors began to take efficiency measures seriously. These two forces led to large-scale retrenchments. Fan (2000) estimates an accumulation of 15 million state sector workers being laid off in 2000, whereas Appleton *et al*. (2001) estimate the number to be 27.7 million. The urban household surveys of 1995 and 1999 conducted by the Institute of Economics, Chinese Academy of Social Sciences, reveal that the urban unemployment rate defined to include laid off workers increased from 8 percent to 17 percent over this period.

Such significant economic restructuring, leading to large increases in unemployment may widen the income distribution in a number of important ways. Initially most of the unemployed who lose employment income may be pushed into the lower end of the income distribution and the income distribution may widen. As time passes high unemployment is likely to have other effects. Wage levels among the unskilled who compete with the unemployed may fall leading to a further widening of the income distribution generated by a widening wage distribution among those with jobs. In addition, the ability of the unemployed to re-enter the work force under conditions similar to the previous income levels may vary considerably and may have long-term income distribution effects. For example, young, educated, and energetic workers who are laid off may easily find better paid jobs elsewhere in the economy or set up their own businesses, while laid off workers who are older or less skilled may fall into long term unemployment or accept jobs in the lower paid informal sector.

This study uses three Household Income Distribution Survey data to investigate the change in income distribution over the gradual and radical reform periods. The surveys were conducted by the Institute of Economics, Chinese Academy of Social Sciences for the years 1988, 1995 and 1999. The questionnaires were designed in a relatively consistent manner for the three years indicated and provide a good basis for a comparative study. The sample size for the three years is 8,992, 6,930, and 4,493 households, respectively. For 1988 the survey mainly covers ten provinces, including Beijing, Shanxi, Jiangsu, Liaoning, Anhui, Henan, Hubei, Guangdong, Yunnan, and Gansu.[2] In 1995, Sichuan province is added to the previous ten provinces. The six provinces included in the 1999 survey are Beijing, Jiangsu, Liaoning, Henan, Sichuan, and Gansu. Thus, only five provinces were surveyed in all three years. For consistency in an over-time comparison, this paper only uses the data from these five provinces. Excluding missing values, the number of households included in this study is 3,700, 2,746, and 3,215 for the three survey years, respectively.[3] The issue of how representative of the five province data

[2]Although there are four households being coded as located in Sichuan province in 1988 data, this is due to mis-coding of the data as Sichuan province is not included in the 1988 survey.

[3]The empirical works conducted later on in this paper are not sensitive to the number of sample provinces included. For each calculation and estimation, the full sample, the sample of six provinces (five consistent provinces plus Sichuan) and five consistent provinces are used and the results obtained from these calculation and estimation are consistent. Although only results from the consistent five province sample are reported other results are available upon request from the author.

TABLE 1

VARIOUS INEQUALITY MEASURES OF INCOME, 1988, 1995, AND 1999

	Real Per Capita HH Income			Real HH Income		
	1988	1995	1999	1988	1995	1999
Relative mean Dv.	0.149	0.194	0.222	0.148	0.184	0.221
Coeff. Var.	0.447	0.558	0.636	0.442	0.514	0.663
SD of logs	0.392	0.506	0.601	0.404	0.484	0.605
Gini coefficient	0.215	0.273	0.313	0.215	0.260	0.313
Mehran measure	0.300	0.380	0.429	0.302	0.364	0.428
Piesch measure	0.173	0.220	0.255	0.172	0.209	0.255
Kakwani measure	0.044	0.068	0.088	0.044	0.062	0.088
Theil entropy measure	0.082	0.126	0.166	0.082	0.114	0.170
Theil mean log Dv.	0.079	0.126	0.171	0.081	0.115	0.173

naturally arises. Using State Statistical Bureau's (SSB) Urban Household Survey Data it is found that the inequality measures presented using five province data are consistent with those using the data for urban economy as a whole for 1995 and 1999.[4] Thus, it is fair to say that the trend of change in inequality found in this study is sufficiently representative of the national trend.

The income variables used are "real household disposable income" and "real household per capita disposable income." The surveys ask direct questions on income (including individual's labor market earnings, government subsidies, income in kind, interest-dividends-rent, and public and private transfers) and income tax of each household member. Household disposable income in this study is defined as the sum of individual income from all sources and household members minus their income tax, plus household incomes which are not attributable to individuals, including from family enterprises and property. Real income is derived by deflating with an urban CPI (1988 = 100) provided by the China State Statistical Bureau (SSB, 2000). The summary statistics of the data are reported in Appendix A.

According to the survey data, average real household per capita disposable income grew from 1,372 yuan in 1988 to 2,101 yuan in 1995, and further to 2694 yuan in 1999. The average annual growth rate is 7.6 percent for the period 1988–95 and 6.4 percent for the period 1995–99. These growth rates are slightly lower than those reported earlier using the national statistical data.

Table 1 presents the measures of income inequality. It is clearly shown that income inequality increased during the period of interest, regardless of the inequality measure or income measure. Using the Gini coefficient as an example, our estimates of the Gini for per capita household disposable income increased from 21.5 in 1988, to 27.3 in 1995, and further increased to 31.3 in 1999. The observed Gini coefficient for 1988 and 1995 are very close to those calculated in Gustafsson and Li (2001), where from the same surveys using ten provinces data they obtained the Gini coefficient for urban per capita income changing from 23.93

[4]This will be discussed in detail later. The data are provided by the Urban Household Survey Team of the SSB.

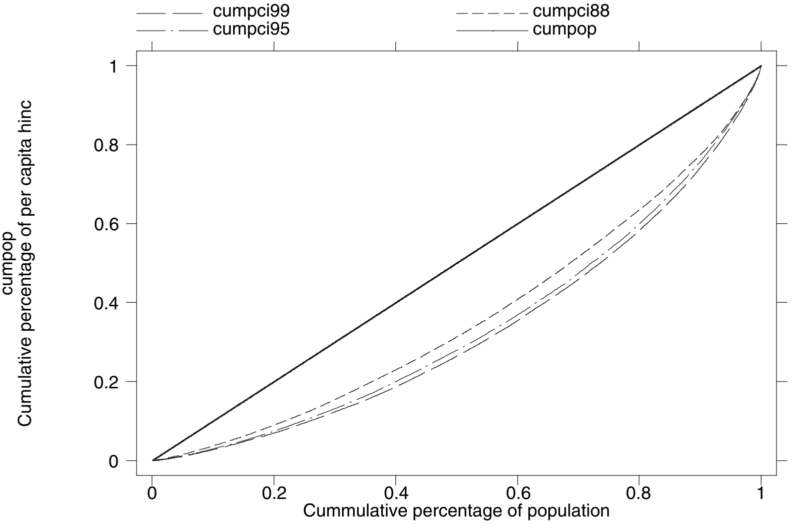

Figure 1. Lorenz Curves for Real Per Capita Household Disposable Income, 1988, 1995, and 1999

in 1988 to 27.55 in 1995.[5] According to the State Statistical Bureau, the Gini coefficients for the urban economy as a whole changed from 22.3 percent in 1988, to 28.9 percent in 1995, and further increased to 30.5 percent in 1999. These data further confirm that the data set used in this study is valid in representing the change in inequality in urban China.

Figure 1 plots the Lorenz curves for the three survey years. The short dashed curve indicates income distribution in 1988, the long dash and dotted line indicates the 1995 situation, and the long dashed curve presents the 1999 situation. If one Lorenz curve lies everywhere above another it is said to "Lorenz dominate" the other curve and all inequality measures will show inequality to be lower for the higher curve. Figure 1 indicates that the 1988 Lorenz curve dominates that of 1995, which in turn dominates that of 1999. This again confirms that income inequality increased over the period.

To understand whether the increase in income inequality affected social welfare over the period studied, Figure 2 presents the generalized Lorenz curve for 1988, 1995 and 1999 (left hand side) and the difference between the curves of 1988 and 1995 and the curves of 1995 and 1999 (right hand side). The vertical axis rep-

[5]Gustafsson and Li (1999) and Khan and Riskin (2001) used the same survey data from the 1988 and 1995 surveys. The disposable income used in their study, however, is adjusted for the housing subsidies received by households who were renting from the government and rental value of owner occupied housing. The way they conducted such imputations are not discussed in details in their works. To check the sensitivity of using our income measure rather than using rental value adjusted income measure in inequality analysis, the simple explanation given in Gustafsson and Li (1999) is followed to carry out the imputation of the housing subsidy and the rental value of owner occupied housing. These imputed values are then added to household income. It is found that adjusting for rental subsidy and rental value of owner occupied housing has very small impact on the measure of income inequality. Some inequality indices using our income measure (original income) and that adjusting for rental subsidy and rental value of owner occupied housing, together with detailed imputation method, are reported in Appendix B.

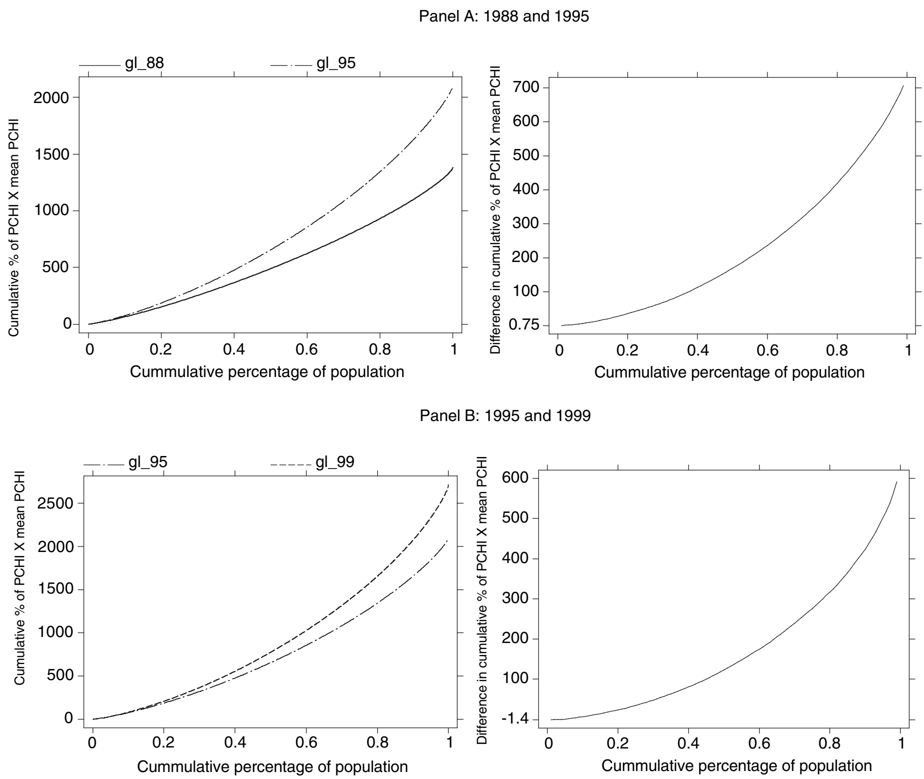

Figure 2. Generalized Lorenz Curve

resents the cumulative percentage of per capita household income multiplied by the mean. It shows the total resources being accessed by each percentile of the population. If one generalized Lorenz curve lies everywhere above another, it is said that the higher curve is preferable to the lower curve with regard to social welfare as every percentile of the population distribution has access to more resources.

Panel A of Figure 2 illustrates the comparison between the 1988 and 1995 distributions. It is clear that although income inequality had increased by 1995, the income growth over the period more than compensated for the inequality increase, as social welfare in 1995 is everywhere greater than in 1988.

Panel B of Figure 2 demonstrates the comparison between 1995 and 1999. The 1999 generalized Lorenz curve lies below that of 1995 at the bottom 5 percent of the distribution (shown by the negative value of the vertical axis of the right hand graph indicating the difference between the two years) and above the 1995 curve everywhere after the bottom 5 percent of the distribution. This result suggests that the social welfare of the bottom end of the income group was worse in 1999 than in 1995, while each percentile above the fifth percentile was better off.

Why, then, did income at the lower end of the distribution fall while medium and high-income families enjoyed significant income gains over the period 1995 to 1999? As economic restructuring may be an important cause, Figure 3 presents the distribution of households with unemployed members across different income

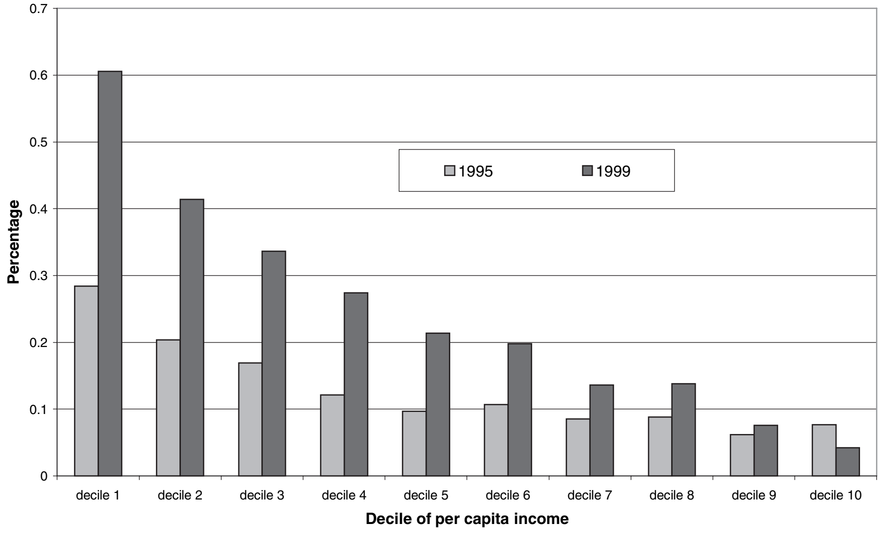

Figure 3. Distribution of Households with Unemployed Members Across Income Deciles

deciles in 1995 and 1999.[6] It indicates that the number of households with unemployed members more than doubled for the lower two deciles while it hardly changed for the top two deciles.

If unemployment is an important cause of the reduction in income at the lower end of the distribution between 1995 and 1999, why is it that not all unemployed households fall into the lowest income group? Perhaps the reduction in household income from one member being unemployed can be offset by income earned by other employed members. Of course, households with more unemployed members are less likely to be able to compensate within the household and hence more likely to fall into the lower end of the distribution. Indeed, in 1999 around 50 percent of the households with two or more unemployed members were located at the bottom decile of the income distribution, and about 30 percent of these households were concentrated at the lowest 5 percentiles of the distribution. In 1995 only 25 percent of households with more than one member unemployed were located in the bottom ten percentiles of income distribution, indicating a better ability for households to compensate for unemployment across household members in 1995 than in 1999.

To sum up, income inequality in urban China has increased. The change, however, may be due to different factors over different periods of economic reform and hence may have different effects on social and political stability.

3. Methodology

Decomposition of the contributing factors to income inequality has long been an important methodological issue in the income distribution literature (Cowell, 2000). Most previous studies employ the decomposition methodology formally

[6]Unemployment in this study is defined as unemployed plus laid off workers.

developed by Shorrocks (1984) and extended by Cowell and Jenkins (1995) and Jenkins (1995). These methods decompose certain inequality indices into between and within mutually exclusive population subgroups. Although this type of decomposition is widely used, it has some shortcomings. One is that the relative contributions of factors critically depend on the order in which the factors are introduced into the analysis (Fields, 1998). In addition, the decompositions require a partition of the population and if the sample size is not large enough, there may not be a meaningful measure of within group inequality for each cell. Given that most empirical studies have limited sample size, the number of contributing factors to income inequality, which can be controlled for, is limited.

Another set of decomposition methodologies have been developed by DiNardo *et al*. (1996) and modified by Cameron (2000).[7] These methods use simulation techniques to impose a structure of income generating factors from the terminal year on the initial year to analyze the contributing factors to the change in income/wage inequality. Although these decompositions are useful for looking at changes in inequality, they suffer from similar shortcomings to Shorrocks (1982) and Cowell and Jenkins (1995).

Recently, regression-based decomposition approaches have been developed (Fields, 1998; Morduch and Sicular, 2002). Assuming an income generating functions $\ln Y_{it} = \sum_j \alpha_{jt} Z_{ijt}$, where j is the number of explanatory variables, Fields (1998) shows that the contribution of an income-generating factor Z_j to the total income inequality can be written as:

$$p_j(\ln Y) = s_j(\ln Y) / R^2(\ln Y) \tag{1}$$

where:

$$s_j(\ln Y) = \operatorname{cov}[a_j Z_j, \ln Y] / \sigma^2(\ln Y) = \frac{a_j * \sigma(Z_j) * \operatorname{cor}[Z_j, \ln Y]}{\sigma(\ln Y)}. \tag{2}$$

The contribution of the income generating factor Z_j to the change in total income inequality measured by any inequality index $I(.)$ over time or between countries/groups can be written as:

$$\Pi_j(I(.)) = \frac{s_{j,2} * I_2 - s_{j,1} * I_1}{I_2 - I_1} \tag{3}$$

where the subscript 1 and 2 refer to time, country, or group.

Fields (1998) points to two important qualifying issues related to the methodology. First, unless there is Lorenz-dominance, it is impossible to determine whether inequality has increased or decreased over time as there always exists an inequality measure that registers an increase and another that registers a decrease in inequality (Sen, 1973). Thus, with Lorenz-crossing it is impossible to specify which factor is responsible for the change in inequality over time. The fact that Figure 2 shows no Lorenz-crossing across the three data periods enables an unambiguous study of the factors contributing to the increase in income inequality.

[7]Bourguignon *et al*. (1998) independently developed a micro-simulation decomposition methodology, which is very similar to DiNardo *et al*. (1996) and Cameron (2000).

Second, the proportion of each factor contributing to the change in income inequality over time depends on the choice of inequality measure used.

Although simple, the Fields' (1998) methodology is particularly powerful. It can identify the factors contributing to the level of income inequality and its change over time (or between countries/regions) without any limitation on the number of factors that can be included in the analysis. In addition, all factors can be controlled for at the same time, hence the results do not rely on the order in which the factors are introduced.[8]

To utilize Fields' (1998) decomposition approach the income generating function needs to be specified. Household income (or per capita household income) is normally defined as a function of the earnings of household members, income from household business, and household demographic characteristics. In the case of urban China, only a limited number of households have a family business, hence income from this source is not important. Thus, household income can mainly be attributed to factors that affect the earnings of household members and demographic features of the household. The household income generating function may be specified as follows.

$$\ln Y_i = \beta_j X_{ij} + \gamma_j R_{ij} + \delta HH_i + \lambda Region_i + \varepsilon_i \tag{4}$$

where $\ln Y_i$ is the logarithmic of household real per capita disposable income (PCHDI) for household i. X_{ij} is a vector of human capital and other factors which affect the earnings of member j in household i, including age and its squared term, years of schooling, and a dummy variable indicating whether or not the individual is a party member. R_{ij} is a vector of factors indicating the effect of economic restructuring on the earnings of individual j of household i. Variables included in this vector contain a dummy variable indicating whether or not the individual is working in a loss making firm, whether or not the individual is unemployed, and the ownership of the enterprise in which the individual is employed. HH_i is a vector of household demographic composition variables, including the gender of the household head, household composition, household size, and the proportion of household members who are in the labor force. Finally, $Region_i$ is the regional indicator of the household.

4. Identifying Contributing Factors to the Change in Income Inequality

4.1. *Determinants of Income Variation*

The income generating model specified in equation (4) is estimated for the three cross-sectional data sets. Table 2 reports the results with the log of real per capita household disposable income as the dependent variable.[9]

Before discussing the results some data and estimation issues need to be addressed. Due to a high multicollinearity between the age and years of schooling of husbands and wives the average age and years of schooling of household

[8]Morduch and Sicular (2002) pointed out some limitations of Fields' decomposition method.

[9]Equation (4) is also estimated using the full sample for each year (10 provinces for 1988, 11 for 1995, and 6 for 1999), the six province sample (five consistent provinces plus Sichuan province), and using log real household income as the dependent variable and the results are consistent. These results are available upon request from the author.

TABLE 2

Determinants of Real Per Capita Household Disposable Income, 1988, 1995, and 1999

	1988		1995		1999	
	Coef.	T-Ratio	Coef.	T-Ratio	Coef.	T-Ratio
Constant	7.6732	73.78	7.1949	42.26	7.9629	39.71
Average age of HH labor	–0.0125	–2.40	0.0333	4.34	0.0041	0.46
(Average age of HH labor)^2	0.0002	3.14	–0.0003	–3.65	0.0000	–0.15
Average years of schooling of HH L	0.0174	8.62	0.0353	13.20	0.0429	12.17
H party membership	0.0611	5.98	0.0699	4.93	0.1064	6.61
W party membership	0.0423	3.00	0.0513	2.85	0.0914	4.59
H being unemployed			–0.1523	–4.84	–0.2901	–10.43
W being unemployed			–0.0947	–3.48	–0.2396	–10.48
2nd generation being unemp.			–0.0990	–2.62	–0.1703	–4.19
H working in loss-making firm			–0.0928	–5.26	–0.1807	–9.85
W working in loss-making firm			–0.0599	–3.41	–0.0876	–4.75
H working in local SOEs	–0.0280	–2.12	–0.1015	–5.91	–0.0669	–3.42
H working in collectives	–0.0740	–4.47	–0.1708	–6.73	–0.1365	–4.42
H working in private sector	0.0496	0.97	–0.0630	–1.25	0.0198	0.58
H did not report sector	0.0891	3.43	0.0582	1.13	–0.2020	–2.41
W working in local SOEs	0.0011	0.08	–0.0302	–1.55	–0.0176	–0.82
W working in collectives	–0.0627	–4.25	–0.0855	–3.77	–0.1121	–4.28
W working in private sector	–0.0379	–1.06	–0.0867	–1.74	–0.0613	–1.77
W did not report sector	–0.0091	–0.54	–0.2049	–6.36	–0.2608	–7.10
Male as the household head	0.0029	0.13	–0.0419	–2.79	–0.0344	–2.03
% of kids aged 0–5	–0.4603	–9.41	–0.1977	–2.24	–0.3669	–3.47
% of kids aged 6–10	–0.3086	–7.13	–0.0007	–0.01	–0.1027	–1.30
% of kids aged 11–16	–0.3903	–8.83	–0.0534	–0.90	–0.2164	–3.34
% of elderly	0.1167	2.02	0.2424	4.21	0.1805	2.85
Household size	–0.2134	–32.75	–0.2720	–21.34	–0.1791	–11.78
Number of employed HH members	0.1963	19.57	0.1879	12.38	0.1345	8.28
Liaoning	–0.0506	–2.84	–0.3115	–14.67	–0.4913	–20.70
Jiangsui	–0.0232	–1.31	–0.0515	–2.43	–0.1647	–6.86
Henan	–0.2955	–17.38	–0.4626	–20.92	–0.5322	–22.68
Gansu	–0.1408	–7.39	–0.4842	–19.54	–0.5114	–20.68
Number of observations	3,700	2,746	3,215			
Adjusted R^2	0.46	0.56	0.52			
R^2	0.46	0.56	0.52			

Note: The central state owned enterprises sector and Beijing are used as the omitted category for the sector of employment and region, respectively.

members who are in the labor force are used. Another issue is that the concept of "working in a loss-making firm" did not exist in 1988 and hence the variable representing it is not included in the 1988 estimation. In addition, the unemployment is very low in 1988 and the variables measure husbands, wives, or the second generation unemployment have to be excluded from the 1988 estimation.

Table 2 reveals interesting results, especially when compared across the three survey years. The discussion below follows the order of human capital related factors, economic restructuring, household composition, and regional effects.

The effect of education on real PCHDI has increased over time. One more year of education increased real PCHDI by 1.7 percent in 1988, 3.5 percent in 1995, and 4.3 percent in 1999. The increase in the return to education reflects the effect of market oriented economic reform in the urban labor market.

The effect of average age of household members who are in the labor force does not appear to have a consistent pattern. Evaluated at the mean age, the effect increased from 1988 to 1995 and then decreased from 1995 to 1999. The change in the patterns of age-income profile could be a result of different factors, such as the change in the shape of the age-earnings profile, family composition, and the macroeconomic conditions of the survey years.

Party members earn significantly higher earnings than non-party members[10] and this effect has increased significantly over the period of the study. In 1988, a household where both husband and wife were party members received 10 percent more per capita income than households without any party member. This ratio increased to 12 percent in 1995, and 20 percent in 1999. This is a significant change, especially from 1995 to 1999 with an increase in return of 8 percentage points. The effect of party membership could reflect either unobservable human capital of party members or their political power. It may be reasonable to believe that with decentralized labor market institutions, the rate of returns to productivity related unobservable characteristics as captured by party membership has risen significantly. It may, however, also be possible that party members received more favorable treatment in 1999 than previously. At this stage it is impossible to disentangle the two effects.

The most important changes over the period of this study come about in the effect of variables representing economic restructuring. The effects on household income of unemployment and working in a loss making firm have changed considerably, though the change in the return to different sectors of employment have not been particularly significant.

In 1988, there were few unemployed individuals (only 0.5 percent of the total labor force), and hence, the variables indicating unemployment are excluded from the 1988 estimation.

Things had taken a significant turn by 1995 when economic restructuring in the urban state sector began to accelerate. Thus, I observe a significant effect of economic restructuring on real PCHDI. A household with a husband unemployed has a 15.2 percent lower PCHDI than a household without an unemployed husband. The income reduction for households with wives or sons/daughters being unemployed is 9.5 and 9.9 percent, respectively. In addition, working in loss-making firms also reduces income substantially. A household with both husband and wife working in a loss-making firm receives 15.3 percent less income than otherwise, which in income terms is equivalent to the household head being laid off. Working in a loss-making firm was generally a first step towards becoming unemployed and in 1995 most laid off workers received a similar pay to those working in a loss-making firm but not yet laid off.

By 1999, radical urban state sector reform had been in place for about 4–5 years and the effect on household income variation was even more severe than in 1995. Households with unemployed husbands, wives, or sons/daughters receive 29, 24, and 17 percent less income, respectively, than households where these members were employed. These ratios are double to triple those observed in 1995. Such significant change may reflect the change in the general economic environment. In 1995, most laid off workers were still being paid by their previous enterprises. By 1999, many of these enterprises were bankrupted and state sector laid off workers

[10]This is a common finding (see, for example, Knight and Song, 1999).

TABLE 3

PERCENTAGE OF DIFFERENT TYPES OF HOUSEHOLDS WITH UNEMPLOYED MEMBERS OR MEMBERS WORKING IN LOSS-MAKING FIRMS

	1995		1999	
	% of HH	Income Reduction	% of HH	Income Reduction
Total number of households	2,746		3,215	
HH without unemp. member	86.27		75.55	
HH with one member being unemp.	11.58	11.0%	21.28	23.2%
HH with two members being unemp.	1.93	22.0%	3.03	46.5%
HH with three members being unemp.	0.22	33.0%	0.06	69.7%
HH with four members being unemp.	0.00		0.03	96.5%
HH with HB being unemp. only	0.81	15.2%	3.98	29.0%
HH with WF being unemp. only	4.83	9.5%	13.15	24.0%
HH with S/D being unemp. only	2.39	9.9%	3.43	17.0%
HH with H&W being unemp.	0.52	24.7%	2.34	53.0%
HH with H&S/D being unemp.	0.10	25.1%	0.33	46.0%
HH with W&S/D being unemp.	0.36	19.4%	1.51	41.0%
HH with H&W&S/D being unemp.	0.05	34.6%	0.27	70.0%
HH with H work in LMFs	10.52	9.3%	12.78	18.1%
HH with W work in LMFs	11.54	6.0%	16.52	8.8%
HH with H&W work in LMFs	11.76	15.3%	19.35	26.9%
Total HH with members work in LMFs	33.83		48.65	

Note: The real per capita household (HH) income equation is also estimated with a variable indicating the number of household members being unemployed for 1995 and 1999 data. The estimated coefficients are –0.0895 and –0.2376 for 1995 and 1999, respectively. The income reduction for the first panel of this table is calculated according to these estimates.

were re-assigned to the re-employment center and received a minimum living standard which is generally lower than the payment from their previous enterprise.

The loss of income for households with husbands working in loss making firms also increased. In 1995, the per capita income of these households was about 9.3 percent less than households without husbands working in a loss-making firm. By 1999, this ratio had increased to 18 percent.

As indicated at the end of Section 2, the poorest 5 percent of households had the highest proportion of households with more than one unemployed member. It is, therefore, worthwhile investigating in more detail how the incomes of these households have been affected by economic restructuring. Income reductions for households with various combinations of unemployed members and/or working in loss-making firms are presented in Table 3. It shows that the income reduction for households with two unemployed members increased from 22 percent in 1995 to 47 percent in 1999. Other things being equal, if a household had both husband and wife unemployed its real PCHDI was reduced by 25 percent in 1995, and 53 percent in 1999. If an additional member (son/daughter) was also unemployed, real PCHDI was reduced by 35 percent in 1995 and 70 percent in 1999.

Table 3 also presents the proportion of households with a different number of unemployed individuals and the demographic composition of unemployment within households for 1995 and 1999. The majority of households with unemployed members had only one unemployed member in both survey years. In addition, the proportion of households with an unemployed wife is very high, especially in 1999 where it is about three times higher than the number of households with only the husband unemployed (13 percent vs. 4 percent). The proportion of households with two members unemployed increased from 1.9 percent in 1995 to 3.0 percent in 1999. Among them the proportion of households with both unemployed husband and wife increased from 0.52 to 2.34 percent over the two survey years. Only a very limited number of households had more than two members unemployed, even in 1999. The group of households with more than one member unemployed deserves more government attention as they are least likely to be able to cushion the effect of economic restructuring within the households.

Table 3 further indicates that "working in a loss making firm" may be a more widespread phenomenon in comparison to being unemployed. In 1995 around 34 percent of the total households has at least one member working in a loss-making firm. This ratio increased to almost 49 percent in 1999. However, the income reduction from "working in a loss-making firm" was not as severe as being unemployed in 1999.

Turning back to Table 2, the pattern of sector of employment does not have a consistent trend, although the superior position of the central state sector (the omitted category) over local state and private sectors seems to have eroded over time, especially when comparing 1995 with 1999.

There is a particularly interesting effect with regard to the household composition. For both 1995 and 1999, households headed by females had significantly higher income than those headed by males. Given that female-headed households accounted for 27 percent of the total sample used in both years, the effect is non-trivial. This is an uncommon result in studies of the determinants of household income for other countries, although Cameron (2000) found a similar phenomenon in Indonesia. In the case of urban China, this could be because female-headed households are more likely to be less traditional and hence better educated with better jobs. Indeed, 36 percent of wives in female-headed households are managerial or professionals in 1995, while in male-headed households only 26 percent of wives belong to this category. These proportions for 1999 are 31 and 23 percent, respectively.

Other household composition variables indicate that relative to working aged adults, households with a higher proportion of young children had lower incomes, while households with more members older than 65 had higher incomes. Larger households had lower per capita incomes, and households with more laborers had higher incomes. These patterns are consistent over the three survey years. In addition, regional dummy variables are important determinant of income variation in all three survey years.

4.2. *Decomposition of Contributing Factors to the Level of Income Inequality and its Change Over Time*

This sub-section quantifies the degree to which the variables included in the income generating equation account for the level of income inequality and its change overtime. In particular, the interest is to identify the relative importance of the impact of economic restructuring compared to the regional effect on income inequality over the three survey years.

TABLE 4

DECOMPOSITION OF THE LEVEL OF PER CAPITA HOUSEHOLD INCOME INEQUALITY[11]

	1988		1995		1999	
	S_j	$P(S_j)$	S_j	$P(S_j)$	S_j	$P(S_j)$
Restructuring	1.19	2.59	9.17	16.28	**18.93**	**36.17**
Unemployment	0.00	0.00	2.01	3.57	8.06	15.39
Loss-making firm	0.00	0.00	2.06	3.66	6.33	12.10
Sector of emp.	1.19	2.56	5.09	9.05	4.54	8.68
Regional effect	10.07	21.97	**19.09**	**33.90**	16.40	31.33
Party	1.52	3.32	2.08	3.70	3.59	6.86
Human capital	3.47	7.58	8.75	15.53	6.07	11.60
Household composition	**29.57**	**64.54**	17.21	30.56	7.35	14.04
Total explained	45.81		56.31		52.34	
Residual	54.19		43.69		47.66	

Table 4 reports the decomposition of the level of inequality on real per capita household income for the three years. The variables included in the estimated equation (4) (Table 2) are grouped into five contributing factors: (1) "economic restructuring," which includes the effect of household members (husband, wife, and sons/daughters) being unemployed, working in a loss-making firm, and their sector of employment; (2) "regional effect," which is captured by the regional dummy variables; (3) "party membership," which includes the two dummy variables indicating whether the husband or wife is a party member; (4) "human capital effect," which is measured by the average age and years of schooling of household members who are in the labor force; and (5) "household composition," which covers the effect of the gender of the household head, the young and old dependency ratios, family size, and the number of laborers in the household.

Notice from Section 3 that Fields' (1998) approach can only decompose the amount of income inequality explained by the variables included in the income regression, that is the proportion represented by R^2s. The rest of the inequality is due to the residual effect. Fortunately, the income generating equations estimated in this study have strong explanatory power. The adjusted R^2s are 46, 56, and 52 percent for the 1988, 1995, and 1999 data, respectively.

The first columns of each of the three years in Table 4 indicate the proportion of the total inequality of the log per capita household income accounted for by each of the five factors and the residual term. The second column takes the total explained portion (the R^2s) as 100 percent and measures the contribution of each of the five contributing factors as a proportion of the total explained component.

The results indicate that the most important contributing factor to income inequality changed from "household composition" in 1988 to "regional effect" in 1995, and to "economic restructuring effect" in 1999. In 1988 there is hardly any effect on income inequality from economic restructuring. In 1995, about 9 percent of income inequality is due to this effect, of which sector of employment accounted

[11]The exercises of decomposition of inequality reported in this table and the change in inequality over time reported in Table 5 are also applied to the regression results obtained from the full sample estimations and the results, which are available upon request from the author, are consistent with what are reported here.

TABLE 5

DECOMPOSITION OF THE CONTRIBUTING FACTORS TO THE CHANGE IN THE GINI COEFFICIENT, 1988–95 AND 1995–99

With 6 provinces data	1988 Gini = 0.215 $S_j(\ln Y)$	1995 Gini = 0.273 $S_j(\ln Y)$	1999 Gini = 0.313 $S_j(\ln Y)$	Change 1988–95 Π_j(Gini)	Change 1995–99 Π_j(Gini)
Restructuring	0.25	2.51	5.92	38.25	**87.51**
Unemployment	0.00	0.55	2.52	9.33	50.53
Loss-making firm	0.00	0.57	1.98	9.59	36.31
Sector of emp.	0.25	1.40	1.42	19.34	0.67
Regional effect	2.16	5.23	5.13	**51.99**	–2.54
Human capital	0.33	0.57	1.12	4.13	14.17
Party	0.75	2.40	1.90	27.98	–12.72
Household composition	6.36	4.72	2.30	–27.80	–61.95
Residual	11.65	11.97	14.92	5.45	75.54
Total	21.50	27.30	31.30	100.00	100.00

for more than half. The effect of economic restructuring increased to 19 percent of the level of income inequality in 1999, of which the effect of unemployment contributed more than 8 percentage points. This indicates that economic restructuring has played an increasingly significant role in the level of income inequality over the period studied. Other effects which have gained some ground in 1999, relative to 1988 and 1995, are party membership and human capital effects, but these do not play as significant a role as the effect of economic restructuring.

The regional effect has always been significant, but it is more so in 1995 than in 1988 and 1999. This result is consistent with that found in Gustafsson and Li (2001) and Khan and Riskin (2001).[12]

The above analysis has identified the most important contributing factors to the level of income inequality in the three survey years. To what extent do these different factors contribute to the increase in income inequality over the periods? To quantify this the decomposition approach specified in equation (3) is applied to the results presented in Tables 1 and 4.

Table 5 summarizes the results for the change in the Gini coefficients between 1988 to 1995, and 1995 to 1999.[13] It shows that the "regional effect" contributed

[12]An issue arises as to whether using five province data may misrepresent the general picture of the regional effect on inequality for the urban economy as a whole. Due to the lack of data for the 1999 survey it is hard to answer this question. However, the 1988 and 1995 surveys include ten consistent provinces and the results for these two years do not change whether to use ten or five province data. The main contributor for income inequality in 1988 is always household composition and in 1995 it is regional effects. This consistency also applies to the analysis of contributing factor to the *change* in the Gini coefficients over time. Regardless which sample is used, the main contributing factor for the change over the period 1988 to 1995 is always the regional effect. If reducing provinces from our sample generated problems with regard to the finding of changing importance of regional inequality, then this would have been true for the 1988 to 1995 changes as well. The fact that little difference between the two samples is found for the 1988 and 1985 data may indicate that the decomposition results are not sensitive to the number of provinces included.

[13]As noticed in Fields (1998), the decomposition of the change in income inequality over time will differ across different inequality indexes. To ascertain that using different inequality indexes provides consistent qualitative results two other inequality indices provided in Table 1 (standard deviation of logs and Theil entropy index) are decomposed as well. The results are consistent with what is observed here and are available upon request from the author.

52.0 percent of the increase in the Gini coefficient between 1988 and 1995, which is the most important contributing factors to the increase in the Gini coefficient in this period. Economic restructuring contributed around 38.3 percent, of which unemployment accounted for 9.3 percentage points, while household members working in a loss-making firm and sector of employment contributed to 9.6 and 19.3 percentage points, respectively. Another important factor that contributed to the increase in income inequality between 1988 and 1995 is the human capital factor, which accounted for more than one quarter of the increase in the Gini coefficient. This finding is consistent with other studies indicating that the effect of labor market reform increased the rate of return to human capital in the 1990s (see, for example, Knight and Song, 2001; Meng, 2000).

During the period 1995–99 the main contributing factor to the increase in the Gini coefficient is "economic restructuring," which accounted for 87.5 percent of the change while the regional effect contributed to the decline of the Gini coefficient. Of the economic restructuring factor, unemployment contributed 50.5 percent of the increase in the Gini coefficient, while working in a loss-making firm accounted for 36.3 percent. The sector of employment has little effect on the increase in income inequality. Another important effect contributing to the increase in the Gini coefficient during this period was party membership.

In both periods, household composition contributed to a decline in income inequality.

5. Conclusions

Urban income inequality in China has increased considerably over the period of economic transition from a planned to a market oriented economy. This study has investigated this change over two important phases of the economic transition: the initial stage of acceleration of the state sector and urban labor market reforms (1988–95) and the period of radical reform in the state sector and the urban labor market (1995–99). An attempt is made to identify the difference in the type of income inequality and the causes of the increase in inequality over these different phases of the economic transition. The main findings may be summarized as follows.

First, although income inequality increased during each of the two periods, the nature of the increase is different. In the first period everybody was made better off and the increase in inequality was due to the relatively stronger income growth at the top end of the distribution. In the second period, however, households at the lowest 5 percentile income distribution experienced an income reduction, while households at the top end of the distribution enjoyed significant income gains. Thus, the increase in inequality in the first period was compensated for by an unambiguous increase in social welfare at every level of income distribution, whereas this was not the case in the second period. Perhaps this is why the social stability has become more of a concern since the late 1990s.

Second, while the increase in income inequality in the first period was mainly due to the increase in regional income variations, this effect was dominated by the impact of economic restructuring in the second period. The increase in the number of households with unemployed members and the significant increase in income reduction for households with unemployed members contributed more than 49 percent of the increase in the Gini coefficient over the period of 1995–99. In addition, households with members working at a loss-making firm also contributed considerably to the increase in income inequality during this period.

An interesting issue, though, is that not all households with unemployed members fell into the low income group. While around 40 percent of such households had income at or below the 20th percentile, nearly 11 percent of households with unemployed members received income above the 70th percentile. This difference may be closely related to the concentration of unemployed members within each household. Indeed, although only 3.7 percent of households in 1999 had more than one unemployed member, on average these households earned more than 50 percent less real PCHDI than other households, *ceteris paribus*. The significant effect of unemployment concentration on income reduction, and hence, on income inequality suggests that, to a large extent, some of the unemployment effect may have been cushioned by within household income transfers. Households whose members are unable to compensate each other are more likely to suffer from severe income reduction due to economic restructuring. Therefore, the households that deserve more government assistance may be those which have more than one member unemployed.

What remains unclear from the current study is whether those who fall into poverty because of economic restructuring will remain in that position for a long period. To examine this issue the duration of unemployment and its impact on income distribution will have to be investigated.

Another equally important issue is related to the long-term trend of income inequality in urban China. Economic restructuring is a transitory phenomenon. Once it is completed, will income inequality reverse back to a more equal level? The answer to this question is probably no. The reason is two fold. First, the relatively equal distribution of income during the pre-reform era and at the beginning of the reform period is mainly the result of a compressed earnings structure. This structure has changed and is continuing to change towards a system where the earnings gap between skilled and unskilled workers is enlarged to reflect market demand for and supply of skilled and unskilled workers. This is indicated by the finding that over the period 1988–99 human capital related factors have been important contributors to the increase in income inequality.

Second, the current large-scale unemployment in urban China is likely to remain for some time as, in part, it is a reflection of the still inflexible urban labor market. There are jobs in urban China. Currently, at any point in time there are more than 50 million rural migrants working in urban China (West and Zhao, 2000). The laid off state sector workers, however, have been unwilling to take these jobs as they are low paid and have low status. As labor market reform proceeds further and the laid off urban workers are more psychologically adjusted toward the new market environment, more and more urban unemployed will accept low paying jobs. Once that happens, unemployment will decrease but income inequality may remain at the current level or even increase.

Two notes of caution should be born in mind when interpreting the results from this study. The data used only include households with urban registration.

Those migrant workers with rural household registration are excluded. Given that rural migrants on average earn much lower earnings and are not eligible for any government subsidies, an urban income inequality measure which takes into account rural migrant households may be larger than estimates in this study and it should be more so for later years when migration becomes more widespread. In addition, our data cover five provinces. Although the general inequality trend obtained from this study is consistent with the national average, there may still be detailed differences.

Appendix A

Summary statistics of the data are presented in Table A1.

TABLE A1

Summary Statistics of the Data

	1988		1995		1999	
Excl. HH with miss. V.	Mean	SD	Mean	SD	Mean	SD
Real PCHDI	1,372.08	585.99	2,101.22	1,187.85	2,694.00	1,700.16
Log(real PCHDI)	7.15	0.37	7.52	0.50	7.73	0.59
Real HI	4,723.24	1,958.54	6,428.41	3,273.15	8,134.93	5,363.45
Log(RHI)	8.39	0.36	8.66	0.47	8.84	0.59
Av. age of HH labor	37.70	7.88	42.02	8.84	42.44	8.35
(Av. age of HH labor)^2	1,482.98	640.02	1,843.57	806.34	1,871.08	765.37
Av. years of sch. of HH L	10.69	2.47	10.38	2.67	10.75	2.33
H party membership	0.39		0.39		0.41	
W party membership	0.14		0.18		0.19	
H being unemployed	0.01		0.05		0.09	
W being unemployed	0.05		0.07		0.15	
2ng gen. being unemp.	0.01		0.03		0.04	
H working in LMFs			0.22		0.32	
W working in LMFs			0.23		0.36	
H work in local SOEs	0.42		0.53		0.43	
H work in collectives	0.14		0.11		0.08	
H work in private sector	0.01		0.02		0.07	
H did not report sector	0.04		0.02		0.01	
W work in local SOEs	0.30		0.47		0.41	
W work in collectives	0.25		0.21		0.17	
W work in private sector	0.02		0.02		0.07	
W did not report sector	0.14		0.07		0.06	
Male as the HH head	0.95		0.73		0.74	
% of children aged 0–5	0.06	0.12	0.04	0.10	0.03	0.09
% of children aged 6–10	0.08	0.14	0.06	0.12	0.05	0.11
% of children aged 11–16	0.07	0.14	0.08	0.14	0.07	0.13
% of elderly	0.03	0.09	0.05	0.15	0.06	0.16
Household size	3.59	0.98	3.20	0.80	3.10	0.73
No. of employed hh members	2.15	0.70	2.25	0.68	2.18	0.68
Liaoning	0.23		0.24		0.21	
Jiangsui	0.27		0.26		0.20	
Henan	0.25		0.20		0.22	
Gansu	0.14		0.14		0.19	
No. of observations	3,700		2,746		3,215	

Appendix B

Imputation of the Rental Subsidy and Rental Value of Owner Occupied Housing

Gustafsson and Li (2001) state that the way they impute the rental subsidy is to subtract the actual rent from the interviewees' self-estimated market rent of the housing they live in. The way they impute the rental value of owner occupied housing is to assume that the rent is 8 percent of the current market value of the house. Questions on the current market value of the house and the self-estimated market value of the rent are asked in the 1995 survey but not in the 1988 survey. It is not clear how Gustafsson and Li (2001) and Khan and Riskin (2001) imputed these values for the 1988 data.

Fortunately, the two questions from the 1995 survey are also asked in the 1999 survey, and hence, the rental subsidy and the rental value of owner occupied housing for the 1995 and 1999 data can be imputed. However, when the data are studied carefully, a couple of issues arise. First, many households did not report their current housing value even though they own their house and many households who rented from the governments did not report their estimated market rental value. This situation is especially severe in the 1995 data (see Table B1). Second, because the market rental value is the households' own estimate, the data are very noisy. Large variations exist and there are noticeable outliers. It is not clear how Gustafsson and Li (2001) and Khan and Riskin (2001) handled these problems.

To fill in missing data, equations are estimated for current market value of owned houses and the market rental value of rented housing using explanatory variables of living area of the house, whether the house has a central heating system, how many rooms in the house (only available for the 1995 data), and provincial dummy variables. The current market value of owned houses and the market rental values of rented government housing for those who fail to report these data are then predicted using the estimated results.

To handle the second problem, the distributions of the household self-estimated market rental values are plotted and the extreme tails of the data are deleted from the estimation of the market rental equations. In the 1995 survey, 0.02 percent of sample households is excluded and the estimated adjusted R^2 increased from 0.03 percent to 0.38 percent, whereas for the 1999 survey 0.04 percent of the sample households is excluded and the adjusted R^2 increased from 0.02 to 0.43. The results of these estimations are reported in Table B2.

The inequality measures of the two income variables are reported in Table B3. They are: (1) original income variable used in the paper (original income); and (2) income including the imputed rental subsidy and imputed rental value of owner occupied housing using predicted value to replace missing values and excluding outliers from the market rental value equation (adjusted income). The results show that adjusting for the rental values does not change the inequality levels and the trend.

TABLE B1

Missing Values for Household Self-estimated Market Rent, 1995 and 1999

	1995		1999	
	Market Housing Value	Estimated Market Rental Value	Market Housing Value	Estimated Market Rental Value
Total number of households	2,896	3,935	2,862	1,278
Number of households reported	2,308	2,865	2,441	1,192
Missing values	588	1,070	421	86
% of missing values	0.20	0.27	0.15	0.07

TABLE B2

Estimated Housing Value and Rental Value Equations for 1995 and 1999

	Current Market Value of Housing		Estimated Market Rental Value of Rented Housing			
			Including Outliers		Excluding Outliers	
	Coef.	T-Ratio	Coef.	T-Ratio	Coef.	T-Ratio
1995						
Constant	22,671.92	3.82	522.67	4.59	498.33	34.29
Living area (meters)	397.48	11.09	4.80	4.36	1.36	9.83
Central heating	18,616.21	6.23	328.29	4.32	143.89	15.32
No. of rooms	1,171.21	1.91	–4.58	–0.35	0.16	0.11
Regional dummies	Yes		Yes		Yes	
No. of observations	2,487		5,364		5,208	
Adjusted R^2	0.35		0.03		0.38	
1999						
Constant	9,152.71	1.76	–235.75	–1.20	–36.42	–1.83
Living area (meters)	1,301.97	22.94	9.03	3.57	6.46	25.17
Central heating	9,346.06	2.52	279.64	1.89	183.63	12.28
Regional dummies	Yes		Yes		Yes	
No. of observations	2,765		4,203		4,187	
Adjusted R^2	0.36		0.02		0.43	

TABLE B3

Inequality Measures of Different Income Variables

	Per Capita Household Real Income		Total Household Real Income	
	Original Income	Adjusted Income	Original Income	Adjusted Income
1995				
Relative mean deviation	0.194	0.196	0.184	0.182
Coefficient of variation	0.558	0.543	0.514	0.494
Standard deviation of logs	0.506	0.505	0.484	0.475
Gini coefficient	0.273	0.275	0.260	0.257
1999				
Relative mean deviation	0.223	0.232	0.221	0.225
Coefficient of variation	0.631	0.640	0.654	0.636
Standard deviation of logs	0.603	0.607	0.606	0.595
Gini coefficient	0.313	0.324	0.313	0.316

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