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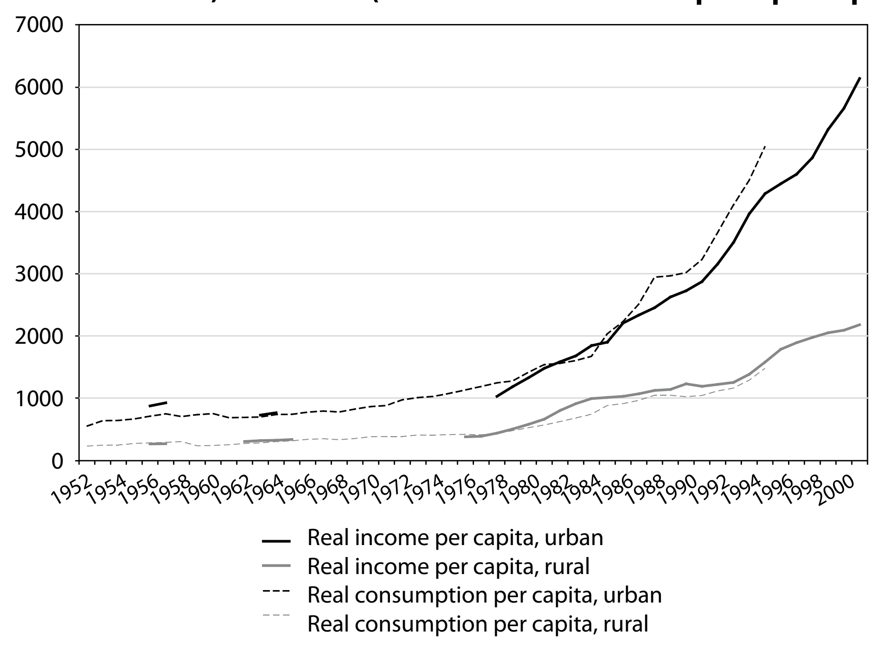
Abstract

In most developing countries, income inequality tends to worsen during initial stages of growth, especially in urban areas. The People's Republic of China (PRC) provides a sharp contrast where income inequality among urban households is lower than that among rural households. In terms of inclusive growth, the existence of income mobility over a longer period of time may mitigate the impacts of widening income inequality measured using cross-sectional data. We explore several ways of measuring income mobility and found considerable income mobility in the PRC, with income mobility lower among rural households than among urban households. When incomes are averaged over 3 years and when adjustments are made for the size and composition of households, income inequality decreases. Social welfare functions are posited that allow for a trade-off between increases in income and increases in income inequality. These suggest strong increases in well-being for urban households in the PRC. In comparison, the corresponding changes in rural households are much smaller.

I. Introduction

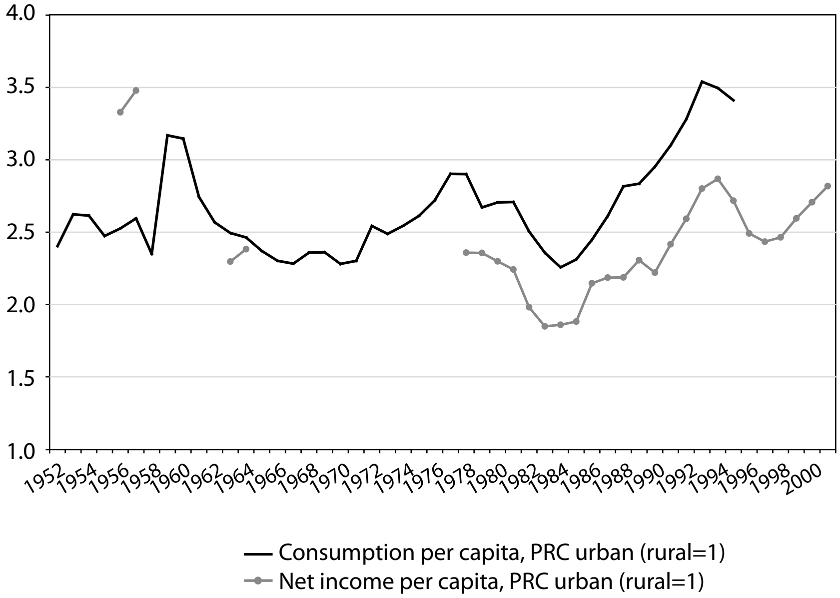
Income inequality remains one of the major challenges of inclusive growth. During the past 50 years, although the income trajectories of both urban and rural sectors in the People's Republic of China (PRC) have been positively trending upward (see Figure 1), urban–rural differences in the PRC have shown a significant increase in the recent years (Figure 2).[1] This increase in income inequality in recent years has been subject to much research and has raised considerable concern (Chen and Ravallion 2004, Khan and Riskin 1998 and 2001, Knight and Song 2003 and 2005, ADB 2007). This paper is concerned with describing and analyzing *longer-run* distribution of incomes in the rural and urban areas of the PRC, and the degree of income mobility that is observed for these households.

Figure 1: Divergence in Urban versus Rural Household Income in the PRC, 1952–2001(real income and consumption per capit



1 In the PRC, the ratio of both net income per capita and net consumption per capita increased sharply after 1984. More specifically, in 1984, the mean urban net income per capita was 2.26 times that of mean rural net income per capita. Ten years later this gap has grown almost 50% so that the urban net income per capita is 3.5 times that of mean rural net income per capita. One observed the same trend using ratios of net consumption per capita.

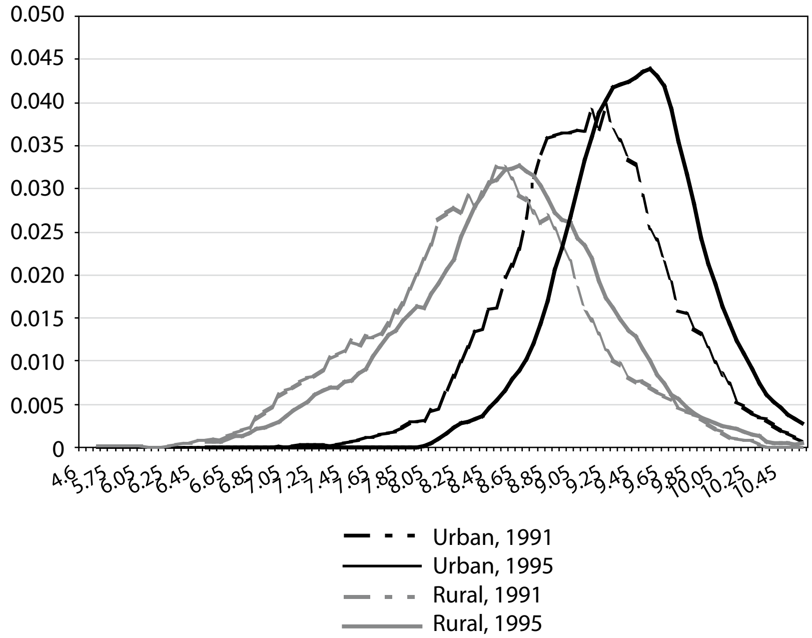
Figure 2: Historical Comparison of Urban–Rural Differences in the PRC (ratio of urban versus rural consumption and net income per capita)



The distinction between the urban and rural sectors of an economy has been a key feature of many models of economic development. Reflecting productivity differences of the activities in the two sectors, the central tendency of rural incomes tends to be lower than that of urban incomes. These income differences form the basis of models of rural–urban labor migration where no restrictions were to be imposed on the movement of labor. Another notable point is that this focus on the central tendency of incomes neglects the fact that income distributions in the two sectors often overlap considerably. Thus, one expression of these urban–rural differences in annual income is provided by Figure 3, which graphs the frequency distribution of household income in the PRC in 1995 among rural and urban households. Clearly, the distribution is displaced to the right among urban households compared with that for rural households. In addition, the income distribution appears narrower for urban households.[2]

[2] The pattern is qualitatively the same in the United States: the rural household income distribution is to the left of that in urban areas. However, the degree of displacement of the rural relative to the urban income distribution in the US is considerably less than that in the PRC (Glaeser and Mare 2001). Data for the PRC are from the Chinese Household Income Project described below. The densities are estimated using the Epanechnikov kernel with a bandwidth of 0.05.

Figure 3: Kernel Density Estimates of the Frequency Distribution of the Logarithm of Household Income in the PRC, 1991 and 1995 (distribution of log of total household income)



In making these comparisons, it is important to recognize that the conventional use of annual incomes may provide a misleading indicator of inequality insofar as one society is characterized by more year-to-year changes in economic status than another society. Hence an important aspect of our analysis is to use the observations on income for the same households to determine the degree to which income inequality in any given year could be smoothed through income mobility over time. In other words, annual income data will provide a misleading indicator of enduring income inequality in societies where there is considerable year-to-year income mobility. How does the rural–urban difference in income inequality based on information on incomes for a single year change when income information over several years for the same households is used to calculate inequality? Are there important differences between rural households and urban households in the degree of income inequality and income mobility?

Government policy could exert a powerful role in shaping the divergent fortunes of urban and rural areas. This is especially true in the PRC, favoring the rural areas in the early years of its move toward a market-based system. Under the "Household Responsibility System" beginning in the late 1970s, agricultural land in rural areas was reallocated rather abruptly to private household plots according to household size (see Walder 2002, Oi 1989). Next came the revival of rural capitalism: development of private household production and marketing of nonagricultural goods and services began in 1979 and accelerated within the next decade or so through the mushrooming of new township and village enterprises (TVEs). The decollectivization, price increases, and relaxation of local trade restrictions resulted in substantial growth in the agricultural economy from 1978 to 1984.

However, agricultural growth decelerated after 1985 and rural families started shifting part of their household labor to off-farm jobs. This would contribute to the rise of inequality we would later observe for rural households in the PRC. Between 1985 and 1995, the output of TVEs in the PRC grew an average of 24% per year in contrast to agricultural growth of 4.2%. TVEs grew to contribute more than 40% of the PRC's industrial output, providing more job options for the rural population. Although as a share of income agriculture still provided the bulk of household income, by the mid-1990s, more than 135 million people in rural areas found off-farm jobs (Nyberg and Rozelle 1999). At the same time, the PRC began reforming its agricultural policies. The nominal rate of protection for four major agricultural commodities decreased significantly from the beginning of reforms in 1978 and continuing through the next two decades. This has different repercussions on the farmers who might have relied on subsidies previously. For the period relevant in our data, this reduction in level of protection and agricultural prices is also accompanied by the intensification of capital use and productivity in the agricultural sector, and the continuing switch from agricultural employment to industry and services.[3] In the late 1980s, government expenditure policy tilted toward industry. Agricultural share of fiscal expenditure did not exceed 15% over the three decades from 1965 to 1996, despite providing the bulk of employment in the PRC. Studies have found that throughout the beginning of the reform and the subsequent two decades, government policies had an anti-agriculture bias. More importantly, the net cash flow from rural to urban areas and from agricultural to industrial sectors was positive and increasing.[4]

The impact of all these government policies on household income mobility would be hard to predict *a priori*. On one hand, government policy has been well known to have an urban bias during the period we examine, which would suggest that accounting for government subsidies would revise urban household mobility upward. Yet, we also see an increasingly large share of rural household income coming from nonfarm sources, which grew from 17% of total household income in 1980 to 47% in 1999 (Anderson, Huang, and Ianchovichina 2004). This could considerably have a positive effect on household income mobility in rural areas.

In our previous study on urban individuals in the PRC during the same period (Khor and Pencavel 2006), the increase in income inequality was accompanied by levels of

[3] In 1990 agriculture accounted for 27% of the GDP and 60% of employment in the PRC. By 1995, the share of agriculture decreased to 20%, and agricultural workers accounted for 52% of all workers.

[4] Chen and Huang (1999) constructed what they call an index of government expenditure bias by dividing the ratio of government expenditure in agriculture to total government expenditure, by the ratio of net income of agriculture to total national income. When investment in agriculture matches its share of national income, the ratio would be 100%, and they defined index values less than that as representing anti-agricultural bias. Between 1965 and 1996, this ratio ranged between 19% (prior the reforms) to 48%.

income mobility higher than observed in the United States (US). A subsequent study found that income mobility significantly decreased for all urban population subgroups in the subsequent few years, while income inequality rose even more (Deng, Li, and Yi 2007). Given the higher degree of within-household risk-sharing in developing countries, we focus on incomes of households in this paper as we investigate the following measurement issues: whether the results are sensitive to the choice of quartiles of household income (quartiles versus clusters), choice of time frame (annual versus multiple years), choice of sample, choice of measures of income (pretransfers versus posttransfers), and implications of measurement errors.[5]

In addition, this paper asks how we evaluate a situation in which incomes are growing at different rates among rural and urban households and, simultaneously, how income inequality is changing. Insofar as society is averse to income inequality, what is the trade-off between increases in income and increases in income inequality? Of course, the answer to this question will depend critically on ethical values, but the economist can provide a representation in which these ethical values are given some quantitative expression. This is our task.

II. Data Sources and Procedures

A. Chinese Household Income Project

We draw information on household income in the PRC from two data sets. The first is the Chinese Household Income Project (CHIP), as discussed in Riskin, Zhao, and Li (2000), which in 1996 surveyed about 8,000 rural households and almost 7,000 urban households.[6] The data are obtained from larger samples designed by the State Statistics Bureau (SSB) though the questions on income differ from SSB's surveys. Nonresponse is unusual although the urban sample excludes those lacking a formal certificate of residence (*hukou*), an exclusion of growing importance as this population grows over time.[7]

5 Our analysis is directed to income inequality, not consumption inequality. We lack successive observations by the same households in the PRC on consumption so the measurement of consumption is not an option for us. If consumption data were available, it would surely be useful to examine them as well as information on income, notwithstanding the well-known problems in imputing the value of services from durable goods and in dealing with commodities infrequently purchased. If they were available, both sources of data are likely to provide insight. This is demonstrated in Knight and Li's (2006) informative analysis of income and consumption from a single cross-section household survey for 1999.

6 The CHIP is a research effort jointly sponsored by the Institute of Economics, Chinese Academy of Social Sciences, Asian Development Bank, and Ford Foundation with additional support provided by the East Asian Institute, Columbia University. Khan and Riskin (2001) provide a careful analysis of some findings.

7 The 2002 survey includes information on those moving to urban areas without a *hukou*. See Deng and Gustafsson (2006) and Ximing et al. (2008).

Table 1: Descriptive Statistics for Households

	CHIP		CHNS*	
Variable	**Urban**	**Rural**	**Urban**	**Rural**
Total income, 1991 (in 1995 yuan)	11111.99	5369.72	5007.45	3279.44
Total income, 1993 (in 1995 yuan)	12753.95	5875.79	6207.08	4681.92
Total income, 1995 (in 1995 yuan)	13743.39	6326.01	7996.88	5646.52
Per adult equivalent, 1991	4531.38	1709.88	1846.75	1150.66
Per adult equivalent, 1993	5184.61	1864.84	2092.61	1663.67
Per adult equivalent, 1995	5600.24	2002.57	2196.58	1728.53
Household size	3.133	4.34	3.379	4.122
Characteristics of Head of Household				
Female	0.339	0.038	0.315	0.092
Age	45.96	44.49	45.96	44.49
Member of Communist Party	0.340	0.149	0.200	0.083
Minority ethnic group	0.042	0.057	0.105	0.161
Average years of schooling	10.42	5.44	8.94	6.08

CHIP = Chinese Household Income Project, CHNS = China Health and Nutrition Survey.

Note: For CHIP, income are expressed in 1995 yuan. For CHNS, the years reported are 1989, 1991, and 1993 and expressed in 1988 yuan in Liaoning.

Measures of income include not only cash payments but also income in kind, state-financed subsidies, and the consumption of agricultural products by households engaged in agricultural production. Here we focus on pretransfer/pretax household income.[8] Though particular results will depend on the concept of income employed, our investigation into the effects of changes in income definitions suggests that our principal findings are robust with respect to alternative definitions of household income. The 1996 survey is based on an earlier survey conducted in the spring of 1989 (Griffin and Zhao 1993, Khan and Riskin 1998) and we compare our measures of income dispersion in 1995 with those in 1988.[9] All income are reported in 1995 yuan, and to mitigate the impact of measurement errors that are most likely to be present in outlying values, we habitually trim the data by omitting the 0.5% of the lowest and the 0.5% of the largest values of income in any sample. Of course, this will reduce the measures of income

8 This is discussed in the Appendix where it is compared with an alternative income concept that incorporates all transfers.

9 The 1988 survey asks about income in a typical month and this is simply converted to annual income by multiplying by 12. In 1995, income is reported as annual income. Details about the formation of our samples from these surveys are outlined in the Appendix. All income information from CHIP is reported in 1995 yuan by applying the consumer price index as a deflator. In their comprehensive analysis of the 1988 and 1995 household income data, Khan and Riskin (2001) use the SSB's consumer price index numbers to deflate rural incomes slightly differently from urban incomes. With 1988=100, the SSB's rural CPI is 220.09 in 1995 and the urban CPI is 227.90 in 1995. They express the suspicion that these price increases understate the amount of inflation over this time. We note the small difference implied in price inflation between rural and urban areas. Benjamin, Brandt, and Giles (2005) compare movements in rural household inequality that deflate incomes with a spatially insensitive price index with those that use a price index that varies across provinces. In any year, the Gini coefficient is some 2–3% lower with the spatially sensitive price index but the movements over time in the Gini coefficient are very similar regardless of the price deflator. Demurger, Fournier, and Li (2005) also compare the effects on inequality indicators of using a provincial price deflator. For urban households in 1995, the Gini coefficient of per adult equivalent household disposable income without such deflation is 0.321 and is 0.298 when a province-sensitive price deflator is used. This difference is similar to that reported for rural households by Benjamin, Brandt, and Giles (2005).

inequality that draw on information throughout the income distribution. When we assessed the impact of this trimming procedure, we found it had inconsequential effects on our important inferences about inequality and mobility.[10]

An important part of this paper consists of the analysis of incomes of the same households over time. From CHIP, the source for this information consists of questions that, in the urban survey, ask respondents to provide their "total income" not only in 1995 but also for each year from 1990 to 1994. For rural households, the retrospective information on income are asked only for the years 1991, 1993, and 1995. Hence in our analysis of this retrospective income information, for urban and rural households together, we are obliged to use data for the above 3 years.

B. China Health and Nutrition Survey

Even though most information about incomes in all surveys is retrospective and even though we have gone to considerable lengths to remove detectable errors, there remains the issue of the extent to which reporting errors in CHIP drive our results. Therefore, after the results using CHIP were derived, we turned to a different source of information, the China Health and Nutrition Survey (CHNS), to assess whether our principal findings for the PRC are replicated in these panel data.[11] The 1991, 1993, and 1997 waves of the CHNS were used to trace the evolution of household income in rural and urban areas of nine provinces in the PRC. As with the CHIP data, the CHNS income data are trimmed by deleting the lowest and highest 0.5% values in any year. CHNS incomes are deflated using a province-specific price index so that all incomes in Liaoning are expressed in 1988 yuan. The principal purpose of using the CHNS data is to provide an independent source of information about the same households over time.

C. Household Size and Composition

In the PRC, urban and rural households tend to be of different size and composition and these differences are not independent of household income. This is suggested by the data in Table 2, which reports the average number of childen (N^C), average number of adults (N^A), and average number of members (N^{A+C}) for each income decile for rural

[10] The measures for households are unweighted by their selection probabilities because the surveys do not supply these. In order to rectify this, we created our own weights using population by provinces and calculated descriptive statistics weighting by the reciprocal of these sampling probabilities. There was little difference between the weighted and unweighted values and, to show this, we report some weighted values in footnotes below. Cowell, Litchfield, and Mercader-Prats (1999) provide an analysis and application of the practice of trimming the tails of income distribution data. The deletion of outliers is a standard (though by no means universal) procedure in labor economics. Card, Lemieux, and Riddell (2004) is a recent example that uses the Current Population Survey, as we do.

[11] We are by no means the first to make use of the household income data in the CHNS. For instance, in a paper that became known to us after the second draft of this paper was completed, Fields and Zhang (2007) make use of both CHIP and CHNS data. Also Benjamin et al. (2008) used the CHNS as repeated cross-sections to describe changes in income inequality from 1991 to 2000. The CHNS is a administered jointly by the Chinese Center for Disease Control and Prevention and the University of North Carolina Population Center (see www.cpc.unc.edu/projects/china.)

and urban households. Rural households tend to be larger than urban households with rural households being more than one-and-a-half times larger than urban households on average. Household size tends to be larger in higher-income households though the link between income and household composition is different between urban and rural areas: the ratio of adults to children tends to be larger in urban areas at higher income levels than in rural areas.

Table 2: Household Size and Composition by Household Income Decile: Rural and Urban Households, 1995 and 1997

	Income Deciles										
	1st	**2nd**	**3rd**	**4th**	**5th**	**6th**	**7th**	**8th**	**9th**	**10th**	**Mean**
PRC, 1995: Rural Households											
N C	1.08	1.11	1.25	1.32	1.38	1.41	1.35	1.34	1.39	1.26	1.29
N A	2.87	2.76	2.72	2.87	2.96	3.08	3.17	3.20	3.27	3.55	3.05
N A + C	3.95	3.87	3.98	4.19	4.34	4.50	4.53	4.54	4.66	4.82	4.34
PRC, 1995: Urban Households											
N C	0.72	0.75	0.79	0.75	0.73	0.72	0.69	0.60	0.56	0.56	0.69
N A	2.10	2.25	2.25	2.32	2.38	2.40	2.47	2.63	2.75	2.92	2.45
N A + C	2.82	3.00	3.04	3.07	3.11	3.12	3.15	3.23	3.31	3.48	3.13
CHNS 1997, Rural Households											
N C	1.30	1.12	1.22	1.20	1.12	1.09	1.11	0.88	0.90	0.71	1.06
N A	2.84	2.98	2.91	2.74	2.74	2.97	2.89	2.68	2.89	2.81	2.85
N A + C	4.21	4.13	4.23	4.03	3.96	4.12	4.07	3.63	3.84	3.56	3.98
CHNS 1997, Urban Households											
N C	0.96	0.69	0.77	0.66	0.77	0.69	0.63	0.59	0.52	0.76	0.70
N A	2.51	2.59	2.70	3.00	2.49	2.92	2.54	2.52	2.48	2.64	2.64
N A + C	3.49	3.35	3.54	3.68	3.30	3.70	3.25	3.18	3.10	3.43	3.41

CHIP = Chinese Household Income Project, CHNS = China Health and Nutrition Survey.

To determine whether our inferences are independent of alternative ways of comparing different types of households, we invoked different adjustments for household size and composition. In addition to using total household income, y_i, with no adjustments for household size and structure, we computed per capita household income, $y_i/(N^A{}_i + N^C{}_i)$, and per equivalent adult household income defined as $y_i/(N^A{}_i + \theta.N^C{}_i)^{\,v}$ where θ is the weight attached to children and v is the scale economies parameter. The implications of alternative values of θ and v were examined and our general inferences did not change noticeably with respect to different values chosen.[12] In the results below, we present per equivalent adult household income using values of θ of 0.75 and of v of 0.85. These values imply that, for example, in evaluating the value of a given yuan or dollar of household income, a household consisting of five adults and no children is "equivalent" to a household with two adults and four children.

[12] We examined values of κ between 0.50 and unity and values of τ between 0.50 and unity. Per capita household income corresponds, of course, to κ = τ = 1. We find that our results, described later in the paper, are not sensitive to the choices of κ and τ.

D. Measures of Income Inequality

To measure income dispersion, in addition to the Gini coefficient, the ratio of income at the 90th percentile to income at the 10th percentile, the coefficient variation of incomes, and the standard deviation of the logarithm of incomes, we present a measure of inequality based on the social welfare function approach to inequality.[13] We draw upon this research explicitly in Section IV below where we assess the change in the well-being in a society when the general level of incomes rises at a time of simultaneously increasing income inequality. For the present, we note the following expression to measure income inequality where *m* denotes the mean of incomes and *n* the number of households:

$$N_{\varepsilon} = 1 - \left[n^{-1} \sum_{i=1}^{n} \left(\frac{y_i}{m} \right)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}} \qquad (1)$$

The computation of this expression requires the specification of the parameter ε: when ε is zero, the index *N* registers indifference to inequality and N_{ε} is zero. But as ε assumes larger values, the index is more sensitive to incomes at the lower tail of the income distribution and *N* increases in value. Common values assumed for ε are between 0.5 and 2.[14]

III. Income Inequality and Mobility among Rural and Urban Households

A. Annual Income Inequality

The first questions to be addressed are the degree of income inequality in urban and rural areas and whether any difference in income inequality measured on the basis of annual incomes is offset by differences in income mobility over time. If household income mobility is different in rural from urban areas, then inequality measured with incomes over a longer period than 1 year may be quite different from inequality measured with annual incomes. To examine this issue, we use first the income information from the 1996 Household Income Project on 5,797 rural households and 6,357 urban households in the PRC with data in all years. A visual representation of the frequency distribution of rural and urban household incomes in 1995 is provided by the kernel densities in Figure 1 from which it is evident that in the PRC, the central tendency of urban incomes is above that of rural incomes. The difference in the logarithm of incomes at the median or the mean implies rural household income is about 43% of urban income.[15]

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13 See, especially, Atkinson (1970) and Blackorby and Donaldson (1978).

14 If ε=1, then $N_{\varepsilon} = 1 - \Pi_i \left(\frac{y_i}{m} \right)^{\frac{1}{n}}$

15 This changes little if familiar differences between urban and rural households are held constant in computing rural–urban income disparity. Thus, holding constant indicators of household size and structure, the age of the

It is also evident from Figure 1 that the annual income distribution among rural households is wider than that among urban households. This visual impression is confirmed by the indicators of income inequality in Table 3. Thus, the Gini coefficient of 1995 household income is 0.354 for rural households and 0.257 for urban households. Whereas incomes at the 90th percentile are about three times that of incomes at the 10th percentile among urban households, they are well over five times among rural households. In general, the indicators of income inequality in urban areas are between one half and three quarters their corresponding values in rural areas. The lower panels of Table 3 indicate that rural income inequality exceeds urban income inequality not only for household income but also for household income adjusted for household size and composition.[16] The CHNS data shows a smaller gap between urban and rural inequality, though income inequality is still higher for rural households.

Table 3: Annual Household Income Inequality: Rural and Urban PRC, 1995 and 1997

	Rural Households		Urban Households	
	CHIP 1995	CHNS 1997	1995	CHNS 1997
Household Incomes				
Gini coefficient	0.354	0.408	0.257	0.357
90th/10th percentile ratio	5.729	7.759	3.203	5.706
coefficient of variation	0.705	0.844	0.495	0.703
standard dev. of log income	0.677	0.827	0.464	0.760
Atkinson's N : ε = 0.5	0.101	0.123	0.053	0.116
Atkinson's N : ε = 1.0	0.196	0.242	0.102	0.234
Atkinson's N : ε = 2.0	0.368	0.460	0.194	0.464
Per Capita Household Income				
Gini coefficient	0.358	0.398	0.265	0.328
90th/10th percentile ratio	5.316	7.403	3.389	5.946
coefficient of variation	0.755	0.822	0.508	0.637
standard dev. of log income	0.677	0.799	0.480	0.694
Atkinson's N : ε = 0.5	0.105	0.129	0.056	0.103
Atkinson's N : ε = 1.0	0.200	0.251	0.109	0.210
Atkinson's N : ε = 2.0	0.373	0.475	0.206	0.439
Per Equivalent Adult Household Income				
Gini coefficient	0.350	0.393	0.254	0.322
90th/10th % ratio	5.200	7.024	3.163	5.553
coefficient of variation	0.721	0.830	0.485	0.629
standard dev. of log income	0.665	0.791	0.459	0.689
Atkinson's N : ε = 0.5	0.100	0.113	0.051	0.094
Atkinson's N : ε = 1.0	0.192	0.224	0.100	0.193
Atkinson's N : ε = 2.0	0.361	0.430	0.190	0.406

CHIP = Chinese Household Income Project, CHNS = China Health and Nutrition Survey.

household head, whether the household head is a Communist Party member, and whether the household head is an ethnic minority, results in mean rural household income being 41% of urban household income. See Khor and Pencavel (2006).

[16] Using a maximum likelihood method to compute an entire distribution from grouped summary information, Wu and Perloff (2005) calculate Gini coefficients of household income of 0.338 among rural households and 0.221 among urban households in 1995, values that are somewhat lower than those in Table 3, albeit the magnitude of the rural–urban difference is similar to the gap we compute. The indicators of income inequality in 1995 among rural households in the PRC in Benjamin, Brandt, and Giles (2005) are slightly lower than those in Table 3. For instance, the Gini coefficient for per capita household income in Table 3 for rural households is 0.358, which is a little larger than the 0.33 they report for their sample of rural households.

B. Indicators of Income Mobility: Income Quintiles

Is there a difference in income mobility between rural households and urban households? A familiar method to address this question is to construct income transition matrices. An income transition matrix cross-classifies households into income quintiles from I (the bottom or poorest quintile) to V (the top or richest quintile) in 2 years. Each quintile contains the same number of households.[17] Each element of the income transition table consists of p_{jk}, the fraction of households in income quintile j in 1 year that occupies income quintile k in a subsequent year.

For the PRC, the years we examine are 1991 and 1995 using CHIP data and 1993 and 1997 using CHNS data. The transition matrix for rural households in CHIP data is presented in Table 4 and the matrix for CHNS households in Table 5 with separate panels for urban, rural, and pooled households. A chi-square test of the null hypothesis that the transition matrices are symmetric cannot be rejected with a high level of confidence.[18] According to the top panel of Table 4, in rural areas, 61% of those who occupied the poorest fifth of households in 1991 were in the same quintile in 1995, whereas in urban areas, 47% of the poorest households in 1991 were still in the lowest income category in 1995. In other words, this particular element of the tables suggests more income mobility in urban than in rural areas. Or consider mobility among the richest households. Among rural households, 60% of those who occupied the richest income quintile in 1991 remained in that same quintile in 1995, whereas among urban households, 53% of those in the top income quintile in 1991 were in the same quintile in 1995. Again, there is a suggestion of greater income mobility in urban than in rural areas. This is also implied by the CHNS data, which show an overall higher level of income mobility than CHIP data, although the rural–urban difference is smaller in these data (see Table 5). The transition matrices based on per capita household income and per equivalent adult household income are similar.

To facilitate comparisons of income mobility, consider three summary indicators of income mobility exhibited in the transition matrices: first, the average quintile move; second, the fraction who remain in the same quintile, also called the "immobility ratio"; and, third, an "adjusted immobility ratio", namely, the fraction who remain in the same quintile plus the fraction who move one quintile.[19] The computed values of these three summary indicators

[17] To ensure an equal number of households in each quintile, if households at the quintile cutoffs have the same income, they are allocated randomly to the adjacent quintiles.

[18] A maximum likelihood test of the symmetry of these transition matrices involves calculating the statistic $\Lambda = \Sigma_{i>j} (p_{ij} - p_{ji})^2 / (p_{ij} + p_{ji})$, which has a chi square distribution with $q(q-1)/2$ degrees of freedom (with q equal to the number of quantiles). For the transition matrices in Tables 4 through 5, the symmetry hypothesis cannot be rejected with a very high level of confidence (i.e., calculated p values close to unity). See Bishop, Fienberg, and Holland (1975, 282–3).

[19] The average quintile move is defined as $\frac{1}{5}\left\{\sum_{j=1}^{5}\sum_{k=1}^{5}\left(|j-k|\right)p_{jk}\right\}$
The fraction who remain in the same quintile is defined as $(5)^{-1}\Sigma_{j=1,..,5}(p_{jj})$. The immobility ratio resembles Shorrocks' (1978) indicator: $(q-T)/(q-1)$ where T is the trace of the matrix and q the number of quantiles (here five).

of income mobility between 1991 and 1995 for rural and urban households are reported at the bottom of Table 4 and Table 5. Within the PRC, income mobility is higher among urban households than among rural households: the average quintile move is higher for urban households and the immobility ratio and the adjusted immobility ratio are lower for urban households compared with rural households.

Table 4: CHIP per Equivalent Adult Household Income Transition Matrix

			1995 I	1995 II	1995 III	1995 IV	1995 V
Rural	**1991**	I	0.613	0.213	0.114	0.035	0.024
		II	0.242	0.361	0.236	0.118	0.043
		III	0.090	0.267	0.311	0.235	0.097
		IV	0.037	0.136	0.251	0.338	0.237
		V	0.017	0.022	0.089	0.274	0.599
Urban	**1991**	I	0.478	0.234	0.157	0.101	0.029
		II	0.294	0.256	0.212	0.157	0.081
		III	0.153	0.249	0.263	0.202	0.133
		IV	0.067	0.206	0.229	0.277	0.221
		V	0.007	0.055	0.139	0.263	0.537
Pooled	**1991**	I	0.702	0.242	0.042	0.012	0.002
		II	0.252	0.445	0.208	0.074	0.021
		III	0.040	0.251	0.360	0.244	0.106
		IV	0.005	0.056	0.303	0.379	0.256
		V	0.001	0.006	0.088	0.291	0.614

Summary	**Rural**	**Urban**	**Pooled**
Average Quintile Move	0.765	0.649	0.600
Immobility Ratio	0.444	0.534	0.500
Adjusted Immobility Ratio	0.835	0.865	0.909

As a reference point, if every entry in the transition matrix (that is, if every value for p_{jk}) were one fifth (sometimes described as "perfect mobility"), the average quintile move would take the value of 1.6, the immobility ratio would be 0.20, and the adjusted immobility ratio would be 0.52. At the other extreme, if the transition matrix were an identity matrix with unit values on the main diagonal and zeros elsewhere (sometimes described as "complete immobility"), the average quintile move would be 0 and the immobility ratio and the adjusted immobility ratio would each be 1. Evidently, the range of values of the average quintile move is from 1.6 to 0, that of the immobility ratio of 0.20 to 1, and that of the adjusted immobility ratio of 0.52 to 1. Higher values of the average quintile move indicate greater mobility and higher values of the immobility ratio and the adjusted immobility ratio indicate less mobility.

Table 5: CHNS per Equivalent Adult Household Income Transition Matrix

		1997 I	1997 II	1997 III	1997 IV	1997 V
	I	0.332	0.253	0.191	0.140	0.084
Rural	II	0.261	0.299	0.210	0.156	0.073
1993	III	0.148	0.218	0.232	0.259	0.143
	IV	0.170	0.162	0.107	0.210	0.262
	V	0.089	0.067	0.170	0.235	0.438

		1997 I	1997 II	1997 III	1997 IV	1997 V
	I	0.345	0.303	0.134	0.143	0.071
Urban	II	0.261	0.239	0.232	0.134	0.136
1993	III	0.246	0.183	0.225	0.211	0.136
	IV	0.070	0.190	0.246	0.232	0.264
	V	0.077	0.085	0.162	0.275	0.393

		1997 I	1997 II	1997 III	1997 IV	1997 V
	I	0.365	0.256	0.197	0.102	0.080
Pooled	II	0.250	0.291	0.227	0.162	0.070
1993	III	0.174	0.195	0.252	0.229	0.150
	IV	0.133	0.180	0.156	0.264	0.267
	V	0.078	0.078	0.168	0.244	0.434

Summary	Rural	Urban	Pooled
Average Quintile Move	1.176	1.178	1.176
Immobility Ratio	0.302	0.287	0.296
Adjusted Immobility Ratio	0.681	0.682	0.677

C. Indicators of Income Mobility: Income Clusters

The indicators of income mobility discussed in the previous paragraphs are not invariant to the extent of income inequality in a society. In other words, a household experiencing a given increase in income is more likely to cross quintiles in an economy with a narrow income distribution than a household experiencing the same income increase in a society with a wide income distribution. Because the inequality of the annual distribution of income is different in rural areas from that in urban areas, consider constructing an income transition matrix defined not on the basis of income quintiles but on the basis of deviations from median income.

To be specific, specify five income clusters as follows: the lowest cluster consists of households with less than 0.65 of the median income; the second cluster consists of households with incomes between 0.65 and 0.95 of the median income; the third income cluster consists of households with incomes between 0.95 and 1.25 of the median income; the fourth cluster consists of households with incomes between 1.25 and 1.55 of the median income; and the fifth cluster consists of households with incomes above 1.55 of the median income. Obviously, if the median is the same in the two societies, the income cutoffs will be the same, but they will correspond to different fractions of households when income dispersion is different in the two societies. In a society with a wide income distribution, more households will be in the income cluster of less than 0.65 of the median compared with a society with a narrow income distribution. Now, however, households experiencing a given absolute increase in income in two societies will be equally likely to cross the thresholds between income clusters.

The consequences for our indicators of income mobility in the PRC of measuring transitions across income clusters rather than transitions across income quintiles are shown in Table 6. There is a tendency for the difference in mobility between rural and urban areas of the PRC to attenuate: as expected, in rural areas where the income distribution is wider, mobility appears to be greater when measured by movements across income clusters than when measured by movements across income quintiles: and, in urban areas where the annual income distribution is narrower, mobility tends to be less when measured by transitions across income clusters than when measured by transitions across income quintiles. However, it remains the case that household income mobility in urban areas in the PRC exceeds that in the rural areas.

Table 6: Income Mobility for the PRC, 1991–1995

	Quintiles		Clusters	
	Rural	**Urban**	**Rural**	**Urban**
Household Income				
Average cluster/quintile move	0.748	0.973	0.847	0.906
Immobility ratio	0.455	0.360	0.473	0.371
Adjusted immobility ratio	0.840	0.744	0.802	0.780
Per Capita Household Income				
Average cluster/quintile move	0.757	0.930	0.824	0.886
Immobility ratio	0.442	0.373	0.466	0.384
Adjusted immobility ratio	0.842	0.763	0.813	0.787
Per Equivalent Adult Household Income				
Average cluster/quintile move	0.765	0.970	0.839	0.913
Immobility ratio	0.444	0.362	0.464	0.367
Adjusted immobility ratio	0.835	0.743	0.801	0.777

D. Income Mobility: Monte Carlo Simulations and Others

One important feature of the mobility measures we use hitherto is that the magnitude and direction of income mobility depends on the ranking of each households with respect to other households included in the sample. To ensure that our results are not the spurious outcome of a singular sample, we conducted nonparametric bootstrap simulations to obtain the distribution of predicted probabilities income mobility. From our original sample of 12,154 households, we pick an independent random set of 10,000 households to derive relative income rankings and the relevant income mobility measures. In Table 7 we report the resulting means and standard deviations over 1,000 replications separately for urban and rural areas. For both urban and rural households, income mobility patterns corroborate our previous observations that income mobility is greater for urban households. Specifically, the probability of staying in the same quintile averages 55% for rural households, and 49% for urban households.

Table 7: Monte Carlo Simulations and Other Robustness Checks of Probability of Upward and Downward Income Mobility

Simulations (N=1,000)	Urban	Rural
Prob(upward)	0.269	0.193
	(0.002)	(0.002)
Prob(no move)	0.486	0.555
	(0.003)	(0.003)
Prob(downward)	0.245	0.252
	(0.002)	(0.002)
Estimated Coefficients: $y_{it} = \alpha_t + \beta y_{i,t-1} + e_{it}$		
□ ^{91,93}	0.829	0.814
	(0.005)	(0.007)
□ ^{93,95}	0.725	0.790
	(0.008)	(0.007)
□ ^{91,95}	0.580	0.716
	(0.009)	(0.008)

Note: The standard deviations are in parentheses. For the Monte Carlo simulations, independent samples consist of 10,000 random households drawn from the pooled urban and rural sample.

We also present another commonly used measure of income mobility, namely the slope coefficient from a regression of current household income (log wave *t+1* income) on lagged household income (log wave *t* income).

$$y_{it} = \alpha_t + \beta . y_{i,t-1} + e_{it} \qquad (2)$$

where the error terms $e_{it} \sim N(0,1)$. When β is zero, income follows a random walk; a value of unity implies complete immobility of income (current household income is completely predetermined by past income). In Table 7, we present these estimated coefficients by urban–rural distinction. Again, the patterns of the estimated coefficients are broadly consistent with our earlier observations: income mobility in urban areas is still greater than that in rural areas (for the whole urban sample, $\hat{\beta}_{urban}$ = 0.580 and for the whole rural sample, $\hat{\beta}_{rural}$ = 0.716 between 1991 and 1995). However, when looking at a 2-year time period, income mobility for urban households between 1991 and 1993 is smaller than that of rural households. Nonetheless, there is a considerable increase in measured income mobility between 1993 and 1995 for both urban and rural households, with a larger decrease in $\hat{\beta}_{urban}$ than $\hat{\beta}_{rural}$. Note that here β does not distinguish between upward and downward mobility.

E. Measurement Errors

What would be the impact of the presence of measurement errors? That measurement error exists is the conclusion of various studies comparing self-reported income against more detailed documentations of income (see Duncan and Hill 1985, Bound and Kruger 1991, Gottschalk and Huynh 2006). If reported income contains measurement errors, then by extension our measures of income inequality and income mobility would be a reflection of not just the distribution of income, but also the joint distribution of these measurement errors and income. Whether measured inequality and mobility would overstate or understate true inequality and mobility would depend upon the distribution of these measurement errors.

With measurement errors, our measured income $y^m{}_t$ is observed as follows:

$$\begin{aligned} y^m{}_t &= y^*{}_t + \nu_t \\ y^m{}_{t-1} &= y^*{}_{t-1} + \mu_{t-1} \end{aligned} \tag{4}$$

where y^* is the true income, while ν_t and μ_{t-1} are measurement errors in this time period and the previous period. The variance of reported income, σ^{2m}, would also contain both the variances of the true income σ^{2*}, the variances of measurement errors $\sigma^2\sigma$, as well as its covariance with true income:

$$\sigma^{2\,m} = \sigma^{2*} + \sigma^2\mu + 2cov(y^*{}_t\,,\,\mu_t) \tag{5}$$

As Gottschalk and Huynh (2006) demonstrated, it is straightforward to see that as long as the covariance between true earnings and measurement error is positive ($cov(y^*{}_t\,,\,\mu_t)>0$), then the presence of measurement error would mean an overstatement of income inequality, given $(\sigma^{2\mu} + 2cov(y^*{}_t\,,\,\mu_t))>0$:

$$\sigma^{2\,m} > \sigma^{2*} \tag{6}$$

On the other hand, the upward bias of measured variance might be tempered if the covariance between true income and measurement errors is negative. To be more specific, measured variance would actually understate true variance if the covariance between measurement error and true income satisfy the following condition:

$$cov(y^*_t\ ,\ \mu_t)/\sigma^{2\mu} < -0.5 \tag{7}$$

The impact of measurement errors on mobility would also depend on the choice of the measure. Consider the following AR(1) income process:

$$y^*_t = \beta\, y^*_{t-1} + \varepsilon_t \tag{8}$$

Substituting equation (4) into equation (8) yields the following:

$$y^m_t = \beta\,(\, y^m_{t-1} - \mu_{t-1}) + \nu_t + \varepsilon_t \tag{9}$$

Using the OLS estimate of β as the measure of mobility and variance as a measure of inequality, classical measurement error would lead to an overstatement of inequality and mobility. However, nonclassical measurement error can lead to potentially offsetting effects and understate both inequality and mobility.

We have not found any studies that describe the measurement error in the PRC data.[20] Given that the measures of income mobility we are focusing on rely on the ranking of incomes, it is not immediately apparent as to which direction the bias of the measurement errors would take. Obviously, if measurement errors were perfectly mean-reverting (i.e., a perfect negative linear function of income), then the introduction of this type of measurement error would be rank-neutral, and would have no impact on our measures of income mobility. Nonetheless, we think that this would be highly improbable. While an extended treatment of the measurement error is not the focus of this study, here we present two simulations of the potential effects of measurement errors upon the mobility indicators for household income per capita in the PRC and in the US.

The first simulation is based upon the assumption of a classical measurement error: the errors are normally distributed with approximately zero mean: $\mu_t \sim N(0,\ \sigma^{2\mu})$. The second simulation assumes a nonclassical type measurement error, and the mean of simulated error approximately 10% of the mean of measured income. Table 8 reports

[20] There have been some studies that describe the characteristics of the measurement errors found in US data. Comparing reported earnings in the Survey of Income and Program Participation to tax records in the Detailed Earnings Record data in 1996, Gottschalk and Huynh (2006) found that measurement errors are mean-reverting and that mean earnings are understated by approximately 15% for the full sample. Bound and Krueger (1991) compared earnings reported in Current Population Survey 1978 data to individual administrative Social Security payroll tax records stretching from 1950 to 1978. They found that measurement errors do not follow the classical assumptions. Instead, the errors are serially correlated over 2 years and are negatively correlated with true earnings (mean-reverting). The measurement errors are found to be distributed in a bell shape, with almost zero means for both men and women, and substantial in magnitude. Bound et al. (1994) report that using the Panel Study of Income Dynamics (PSID), again errors are negatively correlated to true earnings, and ratio of measurement error to total variation of earnings in 1982 and 1986 is 0.151 and 0.302.

the descriptive statistics of these mean and errors. Overall, the introduction of classically distributed measurement errors of the magnitude and variance leads to a mild increase in overall income mobility for urban households in the PRC. Average overall quintile move increases by 3.8% while the immobility ratio decreases by 5.1% (see Table 9).

Table 8: Simulated Measurement Errors for Households in the PRC

	Classical		Nonclassical	
Variable	**1990**	**1995**	**1990**	**1995**
Measurement Error				
Mean	–5.952	–9.790	–344.747	–455.355
Standard deviation	401.641	472.693	300.3761	370.7262
Simulated Income				
Mean	3406.705	4582.547	3745.5	5028.113
Standard deviation	2007.833	2376.749	2179.87	2566.367
Reliability Ratio	0.962	0.962	0.981	0.980

Note: Reliability ratio= Var(true earnings)/[variance(true earnings)+variance(error)].

Table 9: Effect of Measurement Errors on Income Mobility Indicators

	Ratio of New Measures of Income Mobility over the Original Measures* Classical			Ratio of New Measures of Income Mobility over the Original Measures* Nonclassical		
Categories	**(A) Average Move**	**(B) Immobility Ratio**	**(C) Stayers + 1movers**	**(A) Average Move**	**(B) Immobility Ratio**	**(C) Stayers + 1movers**
Characteristics of Head of Household						
women	1.054	0.913	0.981	1.015	0.959	1.003
men	1.043	0.943	0.985	1.004	1.003	0.998
age<30	0.994	0.957	1.031	1.060	0.892	0.959
ag 30-50 years	1.034	0.934	0.988	1.008	0.975	1.000
age > 50 years	1.038	0.986	0.975	1.011	1.000	0.987
edu: < high school	1.059	0.935	0.968	0.997	1.015	0.997
edu: high school+	1.047	0.939	0.973	1.014	0.990	0.984
edu: college+	1.016	0.986	0.998	1.013	0.990	0.983
all	**1.038**	**0.949**	**0.985**	**1.005**	**1.001**	**0.995**

* Ratio of new measures/old measures.

For the second simulation of a nonclassically distributed measurement error, we posit that on average, income is underreported by approximately 10% and that these measurement errors follow a mean-reverting distribution. Nonetheless, these measurement errors increase income mobility measures only mildly for households in the PRC: leading to a 0.5% increase in average quintile move, and an even smaller 0.01% increase in the immobility ratio (see Table 9).

F. Using Alternate Definitions of Income

As an economy transitioning from a planned economy, subsidies traditionally accounted for a large share of household income in the PRC. Although subsidies had been largely phased out by the late 1990s, this is not true for our sample spanning the early 1990s.

Would the inclusion of subsidies change our previous conclusions on the trends of household income inequality and mobility?

In Table 10, we presented an alternative measure of household income based on components defined by Khan and Riskin (1998). For rural households, this augments household income to include the value of self-consumed agricultural products and net transfers. Household income in urban areas includes both net transfers and in-kind housing subsidies. One of the striking results we find is that the median net transfers for rural households is negative.[21] All measures of income inequality increased once transfers are accounted for, although inequality in rural areas is still higher than that among urban households. This widening dispersion is also evident when we pooled urban and rural households together—Gini coefficient increases from 0.355 to 0.386. In other words, at first glance, the net effect of transfers in the PRC appears to be regressive, exacerbating inequality not only within urban–rural areas, but also between urban and rural areas. However, this may well be an accounting illusion since poverty targeting in the PRC during the period was geographically focused, and we are unable to distinguish increases in the incomes of rural household that could be directly attributed to government policy.

Table 10: Comparing Subsidy-Adjusted Inequality Indicators (CHIP)

Inequality Measures of Total Household Income in 1995	Original Indicators in 1995		Using Subsidy-Adjusted Income	
	Rural	Urban	Rural	Urban
Gini Coefficient	0.354	0.257	0.411	0.281
Coefficient Variation	0.705	0.495	1.10	0.563
Standard Deviation	0.677	0.464	0.892	0.519
90th/10th Percentiles	5.729	3.203	5.71	3.506
Total Household Income	**6326.01**	**13743.39**	**8137.99**	**15293.81**

The data presents an additional challenge in trying to address the implication for income mobility, especially since a breakdown of household income is only available in 1995. We restrict this exercise to urban households since most of rural households did not receive these subsidies. Our procedure to impute subsidy levels for income in 1990 is to list additional information on the distribution of in-kind housing subsidies in 1988. We fit a prediction equation for in-kind housing subsidy over the sample of urban households, with characteristics of head of households, household demographics, and province dummies as independent variables. In 1988, in-kind housing subsidy was the largest single category of subsidies received by urban households, amounting to an average of one third of total cash income across all urban households. Food subsidy was on average 10.7% of total household cash income while other types of subsidies would amount to an average of 30%. Including subsidies in household income in 1988 thus effectively doubled the household income on average. As we would see below, each of these types of subsidies is not distributed equally across all households.

[21] The average net transfer for rural households is –50.74 yuan, while for urban households the average is 209.61 yuan, before including in-kind housing subsidies for urban households.

The patterns that we see in these regression results seem to confirm our prior understanding of how subsidies are distributed in the PRC. The sum of total subsidies received by each household varies positively and significantly with the following: income deciles of the households, household size, level of education, age, and Communist Party membership of the head of household. A household in the top income decile receives 2,625 yuan in subsidies more than a household in the bottom decile, and households headed by college graduates and professionals receive more than a thousand yuan in annual subsidies than the households whose head receives less than a primary education. In addition, both the workplace of the head of household and geographical location matter as well. These dependent variables explain approximately 38% in the variation of total subsidies. Given the rank-preserving quality of the distribution of subsidies, perhaps it is no surprise that the inclusion of subsidies, while widening the dispersion of household income, does not materially change income mobility measures (see Table 11).

Table 11: Comparing Subsidy-Adjusted Mobility Measures for Urban Households

	Original Measures			Using Subsidy-Adjusted Income		
Characteristics of Head of Household	(A) Average Move	(B) Immobility Ratio	(C) Stayers + Movers	(A) Average Move	(B) Immobility Ratio	(C) Stayers + Movers
Women	1.015	0.352	0.721	0.931	0.374	0.758
Men	1.042	0.337	0.724	0.973	0.363	0.745
Age<30	1.103	0.352	0.672	1.090	0.331	0.697
dag 30-50 years	1.091	0.330	0.697	1.000	0.360	0.731
Age > 50 years	0.945	0.360	0.763	0.876	0.385	0.785
Edu: < high school	0.970	0.366	0.747	0.939	0.381	0.752
Edu: high school+	1.064	0.334	0.708	0.997	0.362	0.727
Edu: college+	1.095	0.317	0.698	0.993	0.335	0.752
All	**1.033**	**0.343**	**0.722**	**0.954**	**0.373**	**0.749**

Edu = education.

G. Factors Associated with Income Mobility

The indicators of income mobility in Table 6 describe the amount of income mobility across income quintiles over 5 years, but they are silent about those attributes of households that are associated with upward or downward mobility. Moreover, one might think of income mobility as a property that has to be measured not simply between one pair of years but between many pairs of years. Put differently, because there are transitory factors that operate in any given year, the "permanent" probability of upward or downward income mobility is not fully observed using information on only one pair of years. Thus, define π_i as a latent index of permanent income mobility of household i and suppose π_i is a linear function of observed characteristics of the household X_i and unobserved factors, u_i :

$$\pi_i = \beta . X_i + u_I \qquad (10)$$

where u_i is assumed to be distributed normally with zero mean and unit variance. This standardized normal assumption will give rise to the estimation of an ordered probit model.

Although permanent income mobility π i is unobserved, a household's position in the elements of the income transition matrices between 1991 and 1995 in Tables 4 and 5 provides information on the permanent mobility of this household. Based on whether a household occupies an element on the diagonal of an income transition matrix or above the diagonal or below the diagonal, define a new variable z_i with the following features: z_i = 1 for households occupying a cell below the main diagonal (that is, for households experiencing downward mobility), z_i = 2 for households occupying a cell on the main diagonal of the income transition matrix (households experiencing no mobility), and z_i = 3 for households in a cell above the main diagonal of the income transition matrix (households experiencing upward mobility).[22] The relation between the observed variable z_i and the latent variable π_i is given as follows:

$z_i = 1$ if $\pi_i \leq 0$,
$z_i = 2$ if $0 < \pi_i \leq \gamma_1$,
$z_i = 3$ if $\gamma_2 \leq \pi_i$

where γ_1 and γ_2 are censoring parameters to be estimated jointly with β. The X variables consist of household size and the following characteristics of the head of household: gender, age (entered as a quadratic form), years of schooling, an ethnic minority, and, for the PRC, membership in the Communist Party.[23] The implications of the maximum likelihood estimation of the β parameters of equation (10) for the marginal effects are given in Tables 12–13.[24]

[22] Thus, in the income transition matrix in which each element is defined by $\{j, k\}$ where j denotes the income quintile in the initial year and k the income quintile in the final year, z_i = 1 if household i occupies an element where $j > k$, z_i = 2 if household i occupies an element where $j = k$, and z_i = 3 if household i occupies an element where $j < k$.

[23] Age is measured in the year 1995 for CHIP and the year 1997 for CHNS.

[24] Estimated standard errors are in parentheses. For continuous variables, marginal effects are partial derivatives, while for discrete variables, the effects report the consequences of a change in the value of the dummy variable from zero to unity. These effects are evaluated at the mean values of the right-hand side variables. "Age" measures years of age of the head of household. "Household Size" is the total number of adults and children in the household. "Woman" takes the value of unity for a household headed by a woman. "Communist Party" takes the value of unity for a household head who is a member of the Communist Party. "Minority" takes the value of unity for a household head who reports being an ethnic minority. "Years of Schooling" denotes the years of schooling of the household head.

Table 12: Marginal Effects from Maximum Likelihood Estimation: CHIP from 1991 to 1995

	Probability (downward mobility)		Probability (no mobility)		Probability (upward mobility)	
	Rural	Urban	Rural	Urban	Rural	Urban
Woman = 1	0.010	–0.029**	–0.001	0.001	–0.010	0.027*
	(0.026)	(0.011)	(0.002)	(0.001)	(0.025)	(0.011)
Years of	–0.0044**	–0.0116**	0.0001	0.0005**	0.0043**	0.0111**
Schooling	(0.0019)	(0.0017)	(0.0001)	(0.0002)	(0.0018)	(0.0016)
Minority = 1	0.042*	–0.041**	–0.004	–0.001	–0.038**	0.042*
	(0.023)	(0.024)	(0.003)	(0.002)	(0.019)	(0.026)
Communist = 1	–0.011	–0.048**	0.001	0.001	0.011	0.047**
	(0.014)	(0.011)	(0.002)	(0.001)	(0.014)	(0.011)
Years of Age/10	–0.035	–0.200**	0.001	0.008**	0.034	0.192**
	(0.031)	(0.034)	(0.001)	(0.003)	(0.031)	(0.033)
(Age) ^{2}/1,000	0.045	0.199**	–0.001	0.008**	–0.044	–0.190**
	(0.030)	(0.030)	(0.001)	(0.003)	(0.030)	(0.030)
Household Size	–0.017**	0.001	0.001*	–0.001	0.017**	–0.001
	(0.004)	(0.006)	(0.001)	(0.001)	(0.004)	(0.006)

Table 13: Marginal Effects: CHNS from 1993 to 1997

	Probability (downward mobility)		Probability (no mobility)		Probability (upward mobility)	
	Rural	Urban	Rural	Urban	Rural	Urban
Woman = 1	0.041	–0.072**	–0.017	0.008	–0.025	0.064
	(0.036)	(0.042)	(0.034)	(0.041)	(0.035)	(0.043)
Years of	–0.003	–0.003	0.004	–0.002	0.001	0.005**
Schooling	(0.004)	(0.005)	(0.004)	(0.005)	(0.004)	(0.0005)
Minority = 1	0.053*	0.001	0.020	0.019	–0.073**	–0.020
	(0.033)	(0.060)	(0.032)	(0.058)	(0.031)	(0.059)
Commnist = 1	0.054*	–0.004	0.122	0.001	–0.068	0.002
	(0.034)	(0.063)	(0.101)	(0.060)	(0.092)	(0.063)
Years of Age/10	0.042	0.030	0.034	0.031	–0.007	–0.060
	(0.070)	(0.107)	(0.068)	(0.101)	(0.072)	(0.107)
(Age) ^{2}/1,000	–0.018	0.016	0.040	–0.042	–0.022	0.025
	(0.067)	(0.095)	(0.065)	(0.090)	(0.069)	(0.095)
Household Size	–0.023*	–0.024*	0.001	0.015	0.005	0.039**
	(0.012)	(0.014)	(0.008)	(0.013)	(0.008)	(0.014)

In general, the magnitude of the marginal effect of a given variable on the probability of upward mobility is close to the negative of the effect of the same variable on the probability of downward mobility. This is consistent with the symmetry of the income transition matrices, something reported earlier. In the PRC, the marginal effects are not the same in the urban and rural sectors: female-headed households tend to be more upwardly mobile in urban areas than male-headed households whereas no meaningful gender differences in mobility in rural areas are evident.[25] Ethnic minorities tend to be more downwardly mobile in rural areas than nonminorities but such differences are not

[25] Female-headed households in urban areas have a 6% higher probability of upward mobility than male-headed households.

apparent in urban areas. While larger households tend to be more upwardly mobile in rural areas, there is no relation between household size and mobility in urban areas. Though the probability of upward income mobility follows an inverted U-shape with respect to age in both rural and urban areas, it reaches a peak at an age for those about 11 years younger in rural than in urban areas. More years of schooling are associated in the PRC with a greater probability of upward income mobility.[26]

These results indicate the differences in the mobility patterns of rural households and urban households in the PRC. The sharp rural–urban differences in levels of income are exhibited also in rural–urban differences in the factors associated with income mobility. The empirical regularities associated with income mobility among urban households are not the same as the empirical regularities among rural households.

H. A Longer Perspective on Income Inequality

What is the relationship between measures of inequality based on income averaged over 3 years and those based on income in a single year? At least for one measure of inequality, namely, the coefficient of variation of incomes, a precise expression may be derived. Suppose we have observations on incomes for years *r*, *s*, and *t*. Though it is not difficult to generalize the expression below, suppose the income distribution in each of these 3 years is stationary.[27] Then the coefficient of variation of income averaged over the 3 years, *C*, may be written as

$$C = C_r * \left(\frac{1}{3}\right)\left[3 + 2(\rho_{rs} + \rho_{st} + \rho_{rt}\right]^{\frac{1}{2}} \tag{11}$$

where C_r is the coefficient of variation in income in a single year *r* and ρ_{jk} is the correlation coefficient between incomes in years *j* and *k.* Equation (11) expresses the inequality of income averaged over 3 years, *C*, as proportional to income inequality in a single year, C_r, where the factor of proportionality depends on the correlation coefficients in incomes, the values of ρ_{jk}. To help understand equation (11), consider limiting cases. Suppose the correlation coefficients, ρ_{jk}, are all unity, a state of complete income immobility. Then the factor of proportionality is unity and *C* equals C_r. But as the correlation coefficients fall in value, so *C* falls relative to C_r. When all values of ρ_{jk} are zero, *C* is 58% of C_r and it requires negative values of ρ_{jk} to reduce *C* further as a fraction of C_r.

[26] The effects are estimated more precisely in CHIP than in CHNS so the statements in this paragraph hold with more confidence for CHIP than for CHNS.

[27] By stationary, we mean it has the same mean and standard deviation. The assumption of a constant standard deviation, σ, is not egregiously at variance with these data. For instance, for total household income, among urban households, σ for 1993 is 1.10 of σ for 1991 and σ for 1995 is 1.14 of σ for 1991.

In Table 14, we present the values of ρ_{jk} for the PRC. For those correlation coefficients 4 years apart, 1991 and 1995, the correlation coefficients are higher among rural households than among urban households—which is consistent with the earlier result of greater income mobility between 1991 and 1995 in urban areas. Using average values for ρ_{jk} for in equation (11) leads to the suggestion that inequality in the average of 3-year income will be about 93% of income inequality in a single year for the PRC.[28] Indeed, according to the table above, in the PRC, inequality over 3 years of income is between 90% and 95% of inequality measured with incomes for 1995 alone, and this figure is similar in rural and urban areas. The usefulness of equation (11) as a guide to thinking about the effect on measures of inequality of averaging over incomes in a number of years is evident.

Table 14: Correlation Coefficients of Per Adult Equivalent Household Income for the Same Households across Different Years: Rural, Urban, and Pooled Households in the PRC (CHIP 1991–1995 and CHNS 1997–1997)

	CHIP		CHNS	
	1993	**1995**	**1993**	**1997**
Rural				
1991	0.824	0.701	0.453	0.308
1993	1	0.765	1	0.377
Urban				
1991	0.877	0.643	0.338	0.122
1993	1	0.760	1	0.310
Pooled				
1991	0.910	0.768	0.421	0.258
1993	1	0.846	1	

CHIP = Chinese Household Income Project, CHNS = China Health and Nutrition Survey.

What happens to measures of income inequality when we average income over 3 years?[29] Table 15 compares measures of inequality in a single year with measures of inequality using incomes averaged over 3 years. Both CHIP and CHNS data show a consistent decrease in inequality across various indicators once we measure income over a longer term. Again, we observe the recurring pattern of higher inequality in rural areas: Gini coefficients for rural households declined from 0.350 to 0.332 using longer-term income averages, while that for urban households declined from 0.254 to 0.242.

[28] For the PRC, ρ_{rs} is the correlation between incomes in 1991 and 1993, ρ_{st} is the correlation between incomes in 1993 and 1995, and ρ_{rt} is the correlation between incomes in 1991 and 1995.

[29] For CHIP, the single year relates to 1995 and the average of 3 years is formed over the years 1991, 1993, and 1995. For CHNS, the single year describes 1997, and the average of 3 years is formed from incomes in 1991, 1993, and 1997.

Table 15: Per Equivalent Adult Household Income Inequality After Averaging Income Over Years

	Rural Households		Urban Households	
CHIP	**1995**	**1991, 1993, 1995**	**1995**	**1991, 1993, 1995**
Gini coefficient	0.350	0.332	0.254	0.242
90th/10th percentile ratio	5.200	4.666	3.163	3.011
Coefficient of variation	0.721	0.669	0.485	0.461
Standard deviation of log income	0.665	0.625	0.459	0.435
Atkinson's N : ε = 0.5	0.100	0.090	0.051	0.046
Atkinson's N : ε = 1.0	0.192	0.173	0.100	0.091
Atkinson's N : ε = 2.0	0.361	0.327	0.190	0.173
	Rural Households		**Urban Households**	
CHNS	**1997**	**1991, 1993, 1997**	**1997**	**1991, 1993, 1997**
Gini coefficient	0.374	0.301	0.355	0.251
90th/10th percentile ratio	7.024	4.151	5.553	3.673
Coefficient of variation	0.718	0.570	0.627	0.453
Standard deviation of log income	0.750	0.552	0.713	0.484
Atkinson's N : ε = 0.5	0.113	0.071	0.094	0.051
Atkinson's N : ε = 1.0	0.224	0.139	0.193	0.103
Atkinson's N : ε = 2.0	0.430	0.262	0.406	0.208

CHIP = Chinese Household Income Project, CHNS = China Health and Nutrition Survey.

IV. Measures of Changes in Social Well-Being

With higher incomes and yet greater income inequality, as a society, was the PRC better off in 1995 than in 1988? The answer depends, in part, on society's aversion to income inequality: a society that is indifferent to inequality will prefer a situation in which incomes are higher regardless of how these higher incomes are distributed. However, usually, people are not indifferent to increasing inequality, and a given increase in income received by a poor household is regarded as constituting a larger increase in social well-being than the same increase in income enjoyed by a rich household. An indicator of well-being that embodies these relative preferences toward levels of income and their distribution is Atkinson's (1970) additive social welfare function.

$$V = (n)^{-1} \sum_{i=1}^{n} (1-\varepsilon)^{-1} y_i^{1-\varepsilon} \qquad (12)$$

where the parameter $\varepsilon \geq 0$ regulates the trade-off between levels of income, y , and the distribution of income and n denotes the number of households.[30] Using the concept of the

[30] Some intuition for ε may be gained by forming from equation (4) the ratio of the marginal social welfare of an increase in household *j's* income to the marginal social welfare of an increase in household *k's* income:

$$\frac{\partial V / \partial y_j}{\partial V / \partial y_k} \equiv \Delta_{jk} \equiv \left(\frac{y_k}{y_j} \right)^{\varepsilon}$$

Suppose household *k* has twice the income of household *j*. Then giving an extra dollar to household *j* raises social welfare by 2^{ε} times as much as giving an extra dollar to household *k*. With $y_k / y_j = 2$, then $\Delta_{jk} = 4$ if $\varepsilon = 2$; $\Delta_{jk} = 32$ if $\varepsilon = 5$; and $\Delta_{jk} = 1{,}024$ if $\varepsilon = 10$.

equally distributed equivalent income, the measure of inequality implied by this function is given by equation (12) above, allowing us to rewrite V in the more transparent way:

$$V = \left(1-\varepsilon\right)^{-1}\left[\left(1-N_{\varepsilon}\right)\cdot m\right]^{1-\varepsilon} \tag{13}$$

which makes clear the substitution possibilities between income levels as summarized in mean income m and income inequality.[31] From equation (13), solve for the mean income needed to attain a given level of welfare, V, when inequality as measured by equation (1) takes a particular value:

$$m = \left(1-\varepsilon\right)^{1/(1-\varepsilon)} V^{1/(1-\varepsilon)} \left(1-N_{\varepsilon}\right)^{-1} \tag{14}$$

Suppose we observe incomes in two periods, period s and period t, and set social welfare equal to the level enjoyed in period s. Then determine the value of m needed in period t to attain the level of well-being in period s given inequality in period t. Let m^V_t be this constant-welfare level of mean income in period t which, with this expression for social welfare, is given by the simple expression

$$m_t^V = \frac{\left(1-N_{\varepsilon s}\right)}{\left(1-N_{\varepsilon t}\right)}\cdot m_s \tag{15}$$

We may say social welfare has improved in period t over period s if mean income in period t, m_t, exceeds the level of income needed to maintain social welfare constant, m^V_t. In this way, an indicator of social welfare is derived, namely, m_t/m^V_t which allows for trade-offs between increases in the level of incomes and increases in income inequality. When m_t/m^V_t exceeds unity, social welfare in period t has improved relative to welfare in period s. Evidently, m_t / m^V_t depends on inequality in period s and inequality in period t and, because these measures of inequality depend on the inequality aversion parameter ε, this indicator of the change in well-being m_t / m^V_t incorporates society's attitudes toward inequality.

According to the household income surveys for 1989 and 1996, average household income increased in the PRC between 1988 and 1995 by more than 3.4% per year. Table 16 indicates that the increase was considerably larger among urban than among rural households. The increases in per capita household income and in per equivalent adult household income were greater than in household income unadjusted for changes in household size and composition.[32] This is because average household size fell in both

[31] As V is an ordinal representation of preferences, there are no observational consequences from multiplying V in equation (5) by $(1 - \varepsilon)$ and raising the result to the power of $1 / (1 - \varepsilon)$ in which case V is linearly homogeneous in m and $(1 - N\varepsilon)$. When $\varepsilon = 1$, $V = (1 - N\varepsilon).m$ where $N\varepsilon$ for the case where $\varepsilon = 1$ has been defined in the footnote beneath equation (1).

[32] Khan and Riskin (2001) report an annual growth rate between 1988 and 1995 of real per capita household income of 4.48% among urban households (somewhat lower than our value of 5.91) and of 4.71% among rural households (which is higher than our value of 2.86%). As has been emphasized, the sample of households in our empirical work in 1988 and 1995 differs from Khan and Riskin's sample so a difference between their estimates and ours is not surprising. Also, we do not use the same price deflators. Khan and Riskin's growth rate of per capita household

rural and urban areas.[33] As seen earlier, these increases in average household income in the PRC were accompanied by a growth in annual income inequality.[34] In most instances, the increases in income inequality among rural households exceed the increase among urban households. Thus, for a given value of ε, we have an expression that provides an index of the degree to which society was better off (if at all) in the PRC in 1995 with higher incomes that were distributed more unequally in 1995 than in 1988.

Table 16: Percent Annual Average Growth in Mean Real Household Income in the PRC

1988–1995	Rural Households	Urban Households	Rural and Urban Households
Household income	0.775	4.202	3.469
Per capita household income	2.899	5.986	5.804
Per equivalent adult household income	2.377	5.433	5.154

The values of m_t / m^V_t are listed in Table 17 for values of ε between 0 and 2 for rural households, urban households, and for the pooled (urban plus rural) households. When $\varepsilon = 0$, society is indifferent to inequality so, given the increase in income in the PRC between 1988 and 1995, and because the increase in income inequality is disregarded, our welfare indicator should register the largest increase. Indeed, along any row of Table 17, the values of m_t / m^V_t are largest when $\varepsilon = 0$. Thus, for household income, when $\varepsilon = 0$, welfare is 6% higher in 1995 than in 1988 for rural households, 33% higher for urban households, and 27% higher for all households. However, as ε assumes larger values, so the increase in welfare is attenuated because the increase in income inequality between 1988 and 1995 assumes greater importance in our welfare indicator. Indeed, once higher values of ε are posited, there is some doubt that welfare among rural households in the PRC increased between 1988 and 1996. When $\varepsilon = 2$, according to total household income, welfare in urban areas in 1995 was 27% above that in 1988 and welfare for all households in 1995 was 6% above that in 1988; but welfare among rural households in 1995 was some 7% below that in 1988.

income of rural and urban households together is 5.05% compared with ours of 5.64%.

[33] Among rich households, in particular, the fall in average household size is quite remarkable: in the top deciles of annual income, on average among rural households, there was one fewer household member in 1995 than in 1988 and among urban households the drop is almost three quarters.

[34] For their sample of households, Kahn and Riskin (2001) report an increase in the Gini coefficient for per capita household income of from 0.338 in 1988 to 0.416 in 1995 among rural households and from 0.233 in 1988 to 0.332 in 1995 for urban households. The Gini coefficients of household income in 1988 in Wu and Perloff (2005) are 0.300 among rural households and 0.201 among urban households, values close to those in Table 21.

Table 17: Values of Indicators of the Change in Social Well-Being, m_t/m^V_t, in the PRC, 1988–1995

	ε = 0	ε = 0.5	ε = 1.0	ε = 1.5	ε = 2.0
Household Income					
Rural households	1.06	1.03	0.99	0.96	0.93
Urban households	1.33	1.32	1.30	1.28	1.27
Rural and urban households	1.27	1.23	1.18	1.12	1.06
Per Capita Household Income					
Rural households	1.22	1.18	1.14	1.10	1.05
Urban households	1.50	1.48	1.45	1.43	1.41
Rural and urban households	1.48	1.43	1.36	1.27	1.18
Per Equivalent Adult Household Income					
Rural households	1.19	1.15	1.11	1.07	1.02
Urban households	1.47	1.44	1.42	1.40	1.37
Rural and urban households	1.44	1.39	1.32	1.24	1.15

Regardless of the particular value of ε posited, a few general conclusions are evident. For a given ε, well-being rose more for urban households than rural households. This is because, as Table 16 indicates, real incomes increased more in urban areas than in rural areas and, as Table 25 shows, income inequality increased less in urban than in rural areas. For a given value of ε, the increase in welfare is greater for per capita household income and least for total household income. This is because the size of households fell between 1988 and 1995, something not taken into account when using total household income as the basis for computing welfare changes. These declines in the size of households are especially marked among households with higher incomes. The reductions in household size between 1988 and 1995 account for the fact that the welfare indicators suggest greater improvements in well-being using per capita household income or per equivalent adult household income than total household income. The reductions in household size that are greater for rich than poor households account for the fact that the difference between the values of the welfare indicators based on household income and the values of the welfare indicators based on per capita household income (or on per equivalent adult household income) falls as ε rises; that is, the difference between inferences based on total household and inferences based on per capita household income falls as increasing weight is placed on income inequality in the evaluation of welfare.

One may ask a slightly different question. Instead of positing a particular value for ε and then determining how welfare changed, one may ask what value of the inequality-aversion parameter is required for welfare in 1995 to be the same as welfare in 1988. Expressed differently, given the actual changes in incomes, how much aversion to income inequality is required for social welfare not to have increased (that is, for m_t / m^V_t to be unity).[35] The answers to this question are provided in Table 18, which highlights the different experiences of rural households and urban households. Given the income changes among urban households, a substantial inequality aversion of over 10 is needed

[35] This involves calculating the value of ε that satisfies $(1 - N \varepsilon s)/(1 - N \varepsilon t) = m t / m s$.

to avoid the inference that social well-being between 1988 and 1995 did not increase. By contrast, among rural households, on the basis of total household income, an aversion to income inequality ε of a little less than unity is needed for social well-being in rural areas not to have increased. This underlines the sharp difference in the experience of urban and rural households.

Table 18: Values of ε Needed for Social Welfare in the PRC Not to Have Increased between 1988 and 1995

	Rural Households	Urban Households	Rural and Urban Households
Household income	0.90	10.04	2.47
Per capita household income	2.47	43.28	3.08
Per equivalent adult household income	2.20	33.51	2.92

V. Conclusions

It is useful, if not essential, to place the facts of urban–rural income differences in the PRC in a comparative context.[36] This is particularly the case in view of the abiding issue of the degree to which different degrees of income inequality are linked to alternative systems of economic organization. Whether capitalism generates more enduring inequalities has been a major question of social analysis for at least well over 150 years. The facts regarding income inequality in the PRC today would appear to bring some empirical light to this question especially as the country has been moving away from a state-directed planned economy and toward a more decentralized market economy. Our findings point to a situation of increasing annual income inequality combined with income mobility higher than that of some developed countries such as the US, despite restrictions on geographical mobility and the high degree of geographical stratification of income. The variables associated with income mobility are not the same among rural households as among urban households.

How does income mobility in the PRC compare to other nondeveloped countries? In general, we observe that measured income mobility seems to be higher in developing countries than that in developed nations: the percentage of those remaining in the same quintile over 5 years is larger than 50% in most of the developed countries listed in Table 19, compared to approximately 40% in Peru, South Africa, and Viet Nam.

[36] In a previous draft of this work, we compared households in the PRC with those in the US. We find greater annual income inequality and less income mobility in the US—a conclusion that holds when examining the annual incomes of rural households and of urban households separately. When incomes are averaged over 3 years and when adjustments are made for the size and composition of households, income inequality among all households differs little between the PRC and the US in the 1990s. Moreover when pooling rural households and urban households and when measuring annual income inequality and income mobility of the pooled households, the mobility of incomes of households in the US differs little from that in the PRC. To be specific, income mobility appears greater in the US after adjusting these incomes for the size and composition of households. In this way, the clear differences in income inequality and income mobility that are evident on first analysis attenuate and sometimes disappear once the greater diversity in household structure in the US is considered.

Table 19: Comparing Year-to-Year Income Mobility in Other Countries

	Time Period (years)	Years Observed					
		Initial	End	(A)	(B)	(C)	Unit of Observation
PRC(1)	1	1990	1991	0.285	0.735	0.984	urban individuals
PRC (1)	1	1991	1992	0.306	0.718	0.981	urban individuals
PRC (1)	1	1992	1993	0.390	0.663	0.961	urban individuals
PRC (1)	1	1993	1994	0.417	0.643	0.953	urban individuals
PRC (1)	1	1994	1995	0.656	0.493	0.886	urban individuals
US (1)	1	1994	1995	0.412	0.665	0.942	urban individuals
US (1)	1	1995	1996	0.354	0.708	0.954	urban individuals
US (1)	1	1996	1997	0.354	0.713	0.949	urban individuals
US (3)	1	1974	1975	0.476	0.611	0.924	men, 20-58 in 1974
Hungary	1	1992	1993	0.724	0.479	0.857	household income per capita
Hungary	1	1993	1994	0.674	0.504	0.868	household income per capita
Hungary	1	1994	1995	0.572	0.563	0.899	household income per capita
Russia	1	1994	1995	0.959	0.406	0.758	hh inc. per adult equivalent
Russia	1	1995	1996	0.996	0.380	0.748	hh inc. per adult equivalent
PRC (1)	2	1991	1993	0.402	0.680	0.977	Urban, hh. inc. per capita
PRC (1)	2	1993	1995	0.561	0.571	0.954	Urban, hh. inc. per capita
PRC (1)	2	1991	1993	0.449	0.685	0.971	Rural, hh. inc. per capita
PRC (1)	2	1993	1995	0.553	0.614	0.963	Rural, hh. inc. per capita
PRC (1)	2	1991	1993	0.347	0.682	0.974	pooled, hh. inc. per capita
PRC (1)	2	1993	1995	0.454	0.592	0.959	pooled, hh. inc. per capita
PRC (1) decile	2	1991	1993	0.730	0.467	0.865	pooled, hh. inc. per capita
PRC (1) decile	2	1993	1995	0.948	0.365	0.783	pooled, hh. inc. per capita

hh. inc. = household income.
(A) Average quintile move.
(B) Immobility ratio.
(C) Fraction of stayers plus those who moved one quintile.
Sources: Measures of income mobility are computed using transition matrices from the following sources:
PRC (1): own tabulation from CHIP data.
US (1) : own tabulation using PSID data.
US (3): Gottschalk (1997, Table 1).
Hungary: Galasi (1998, Table 4).
Russia: Bogomolova and Tapilina (1999, Tables 1–2).

While it might be instructive to compare the PRC with other transitioning markets, we see that even for markets undergoing transition at similar times, there seems to be a variety of experiences in terms of income mobility.[37] Hungary and Russia transitioned to a market-based economy approximately at similar times (between 1989 and 1991). In 1994, the annual measured average quintile move is 0.572 for Hungary, 0.996 for Russia, and 0.347 for the PRC households in 1993. The comparison itself might not be

[37] Although there exists a debate over the implications of market transitions for income inequality and mobility, we do not have any a priori hypothesis about the directions of these changes. Nee (2000) argues in his "market transition theory" that a move toward a market-oriented economy would imply, among other hypotheses, a shift of bargaining power away from government cadres and planners towards producers. Assuming that initial income distribution favors those with larger bargaining power (cadres), this would lead to an increase in income mobility as well as a decrease in inequality by the reduction in that gap. Countering this idea, Walder (1996) states that the shift to market allocation "has no inherent consequences for the allocation of power and income." Our results do not provide specific answers to the variables proposed by Walder (1996) to quantify and test the implications of these market transition theories. The influences of cadre power, characteristics of enterprises, and path dependence on income mobility would have to be relegated to future research. However, we note here than from 1988 to 1995, economic returns to both schooling and being a Communist Party Member have increased.

very meaningful since the unit of analysis upon which the transition matrices are based vary from country to country, but looking at the trends, we observe that the year-to-year income mobility in the PRC increased from 1990 to 1994, while decreasing in Russia and Hungary. This increasing trend in income mobility for the PRC individuals and households is evident regardless whether we measure income mobility using quintiles or deciles. While the year-to-year mobility in the PRC may not be the highest relative to the other countries, longer-term mobility seems to be relatively high (see Table 20). Thus, income mobility may attenuate the more pernicious effects of increasing annual income inequality that we observe for this period in the PRC.

Table 20: Comparing Income Mobility in Other Countries (4–6 years)

	Time Period (years)	Years Observed					
		Initial	End	(A)	(B)	(C)	Unit of Observation
PRC (1)	4	1991	1995	0.568	0.519	0.920	Pooled household income per capita
PRC (1)	4	1991	1995	0.723	0.486	0.905	Urban household income per capita
PRC (1)	4	1991	1995	0.665	0.556	0.937	Rural household income per capita
PRC (1)	5	1990	1995	1.056	0.334	0.711	urban individual
US (1)	5	1993	1998	0.624	0.522	0.888	urban individual
US (2)	5	1986	1991	0.660	0.514	0.868	urban individual
United Kingdom	5	1986	1991	0.660	0.514	0.868	urban individual
Sweden	5	1986	1991	0.684	0.505	0.866	urban individual
Italy	5	1986	1991	0.685	0.503	0.857	urban individual
Germany	5	1986	1991	0.647	0.523	0.876	urban individual
France	5	1986	1991	0.683	0.530	0.854	urban individual
Denmark	5	1986	1991	0.812	0.462	0.810	urban individual
South Africa	5	1993	1998	0.992	0.384	0.742	urban household income per adult equivalent
PRC (2) (rural)	5	1978	1983	0.754	0.444	0.853	rural household income
PRC (2) (rural)	6	1983	1989	1.066	0.340	0.717	rural household income
PRC (1)	4	1991	1995	0.748	0.455	0.840	rural household income
Peru	5	1985	1990	0.892	0.389	0.799	urban household expenditure per capita
Peru	6	1990	1996	1.051	0.323	0.712	urban household expenditure per capita
Viet Nam	6	1992	1998	0.830	0.405	0.805	household expenditure per capita
PRC (2) (rural)	11	1978	1989	1.207	0.298	0.657	household income
US (3)	17	1974	1991	0.964	0.393	0.760	men, 20–42 years old

(A) average quintile move.
(B) immobility ratio.
(C) fraction of stayers plus those who moved one quintile.
Sources: Measures of income mobility are computed using transition matrices from the following:
- PRC (1): own tabulation from CHIP data.
- US (1) : own tabulation using PSID data.
- US (2), United Kingdom, Sweden, Italy, Germany, France, Denmark : OECD (1996).
- South Africa: Woolard and Klasen (2004, Table A3).
- PRC (2): Nee (1994, Table 3), as reproduced in Fields (2001, Table 7.7).
- Peru: Herrera (1999), as reproduced in Fields (2001, Table 7.1).
- Viet Nam: Glewwe and Nguyen (2002, Table 2).
- US (3): Gottschalk (1997, Table 1).

In both urban and rural areas, household incomes have tended to grow at a time when income inequality has widened. If societies are averse to income inequality, from a social welfare perspective, has the growth in incomes offset the increase in income inequality? The answer requires a judgment about society's values and, in particular, about the weight placed on income inequality in the expression of society's welfare. Using a metric for social welfare that compares actual incomes in 1995 with those incomes needed to maintain well-being the same as in 1988, we find unambiguous increases in social well-being for urban households in the PRC where the strong rise in incomes clearly offsets the relatively small increases in income inequality. Among rural households in the PRC, if changes in household size and composition are neglected, the modest increases in incomes were adequate to compensate for increases in income inequality only when society exhibits low levels of aversion to income inequality.

To check the robustness of the findings on income mobility, we present several alternatives of measuring income mobility apart from indicators calculated using income transition matrices based on income quintiles. First we constructed income clusters using median income as an anchor. Next we performed Monte Carlo simulations to ensure that the results did not arise out of any particular singularity of our sample. Simulations of measurement errors did not show a large impact on the measures of income mobility, neither did the inclusion of net transfers. Correlations of income over the available years and regressions of current income on lagged income also show a higher degree of income mobility in urban areas than for rural households. In addition, we find that the degree of income mobility for households is smaller than that for individuals, which could be indicative of risk-sharing within households.

These results have several implications for policy. First, policies targeted toward inclusive growth ought to take into account longer-term income. A focus on annual income may overstate the degree of income inequality faced by individuals and households. Second, means-tested programs ought to take into consideration not just individual income, but also household income. Third, inclusive policies ought to be designed so as to not exacerbate existing inequality. Surprisingly, we find some evidence that public subsidies and transfers actually widened the dispersion of income in the PRC for the relevant time period. However, further research would be required to ascertain this as fact.

Appendix: Data Selection and Procedures

In CHIP data, the risk that the sample does not fully reflect the PRC population is made more serious because not all households with income information for 1995 are represented with their income data for 1991 and 1993. In part the problem, presumably, is one of nonresponse. Among 7,997 rural households with income data for 1995, there is usable income information in 1991 and 1993 also for 72% of them.[38] Among 6,932 urban households with income information in 1995, income information for 1991 and 1993 is available for 92%.

The problem of missing data poses the same sort of concern as the problem of attrition in panel data: when observations are missing nonrandomly, the sample of households with income information in all years is not representative of all households. To help evaluate this, we may determine if the households with missing income data from CHIP for 1991 and 1993 in the PRC are systematically different from all the households who provided income information in 1995. To this effect, define a variable, *Q*, which takes the value of 1 for a household in the PRC with income information for all years (1991, 1993, and 1995) and 0 otherwise. Express *Q* as a function of a number of variables including the household's income in 1995 to determine whether those households without income information in 1991 and 1993 are drawn randomly from all parts of the 1995 income distribution. The relationship is computed by conventional logistic maximum likelihood methods to estimate *Q*, and Table 4 reports the estimated effects of differences in the predicting variables on the probability of complete income information.[39]

Among both urban and rural households in the PRC, the coefficient estimates attached to the income decile dummy variables suggest that the largest differences are associated with the richest households in 1995: for households in the top income decile in 1995, the probability of providing complete income information is 19% in rural areas and 11% in urban areas below the probability in the lowest income decile. In urban areas, there is also the suggestion that complete income information is almost 3% lower in the lowest income decile in 1995 than in fourth income decile. Therefore, the sample from CHIP providing complete income information doest not appear to be entirely representative of all households with well-off households, in particular, less likely to be included in the income data for all years. Consequently, indicators of income inequality for the year 1995 assume lower values for the sample of households with complete income information than for the entire sample. Because of problems of sample attrition and nonresponse, it is not uncommon for studies of long-run income inequality and income mobility to be conducted on

[38] One problem with the rural CHIP file is a suspiciously large number of zero values for household income. Do these zeros really mean no household income or more likely was the information on income not recorded? In 1995, there are 11 households out of 7,997 with zero household income, there are 1,602 with zero income in 1993, and there are 2,060 with zero reported income in 1991. We have dropped all households reporting zero income from our analysis, and this is why the 7,997 households in 1995 shrinks to 5,797 for our analysis sample (that is, we work with 72% of the 1995 sample). As is well known, zero incomes may induce measurement difficulties for inequality indicators because some indicators are not well-defined or assume their limiting values in the presence of zeros (for example, Atkinson's indictor with ε=1 reaches its maximum value when incomes are zero). Issues concerning the interpretation and management of zero income values in surveys are addressed by Cowell, Litchfield, and Mercader-Prats (1999).

[39] In urban areas, there are 6,932 households with income data in 1995 and there is income information for 1991 and 1993 on 6,357 of them. In rural areas, of the 7,997 households with 1995 income data, there are 5,797 households with income information also in 1991 and 1993.In Table 4, estimated standard errors are in parentheses. For continuous variables, marginal effects are partial derivatives while for discrete variables, the effects are of a change in the value of the dummy variable from 0 to 1. These effects are evaluated at the mean values of the right-hand side variables.

samples of individuals or households that are not fully representative of the larger population. However, the fact that this is a frequent feature of research studies on these topics does not mean we may dismiss the seriousness of the potential problem that our inferences about the PRC will be drawn from a sample not entirely representative of the population.

In addition to the problem of nonresponse, there is the problem of response error. The consequences of such measurement error on our measures of income mobility are difficult to assess without knowing the properties of the errors. Some results in the literature regarding measurement error in income are based on the presumption that measurement errors take the classical form, but measurement error in income is unlikely to be classical (Bound et al. 2001, Hyslop and Imbens 2001, and Gottschalk and Huynh 2006). Perhaps the most probable form of response error is that, independent of their true incomes, individuals report the same income (or the same fraction of income) in different years. If this occurs, this will suggest less change in the income distribution than is really the case, and our measures of income mobility will provide a lower bound on true income mobility.

Appendix Table: Marginal Effects from Logit Estimates of the Probability of Providing Information on Household Income in All Six Years

Variables	Urban	Rural
Age	0.004	0.017
	(0.002)	(0.003)
Age-squared	−0.004	−0.000
	(0.002)	(0.000)
Woman=1	−0.008	−0.038
	(0.007)	(0.026)
Communist Party=1	0.004	0.019
	(0.007)	(0.014)
Ethnic minority=1	0.016	−0.143
	(0.014)	(0.022)
Schooling Levels		
College and above	−0.006	0.129
	(0.015)	(0.097)
Professional School	0.017	0.058
	(0.012)	(0.073)
Technical/Vocational School	−0.005	−0.033
	(0.013)	(0.043)
Upper Secondary School	−0.002	0.015
	(0.012)	(0.017)
Lower Secondary School	−0.031	−0.003
	(0.017)	(0.012)
Income Percentiles		
10–20 percentile	0.021	0.007
	(0.012)	(0.023)
20–30 percentile	0.019	0.007
	(0.012)	(0.023)
30–40 percentile	0.027	−0.003
	(0.012)	(0.023)
40–50 percentile	−0.004	0.008
	(0.015)	(0.023)
50–60 percentile	0.009	0.021
	(0.014)	(0.023)
60–70 percentile	0.013	−0.035
	(0.013)	(0.024)
70–80 percentile	−0.006	−0.043
	(0.011)	(0.024)
80–90 percentile	−0.036	−0.116
	(0.018)	(0.026)
> 90 percentile	−0.114	−0.189
	(0.026)	(0.027)
Number of adults	0.009	−0.001
	(0.004)	(0.005)
Number of children	0.016	0.0048
	(0.007)	(0.005)

Note: There are 6,932 urban households and 7,998 rural households. Estimated standard errors are in parentheses. For continuous variables, marginal effects are partial derivatives while for discrete variables, the effects are of a change in the value of the dummy variable from 0 to 1. These effects are evaluated at the mean values of the right-hand side variables. The number of adults and number of children refer to those residing specifically in the household (with someone 18 years or over constituting an adult). All the other demographic variables are dichotomous variables associated with the characteristics of the head of household. The "x-y percentile" are dichotomous variables that take the value of unity for households with an income in 1995 in the percentile range between x and y. The lowest 10th percentile constitutes the reference category. For schooling, the reference category is those with an elementary education or less.

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About the Paper

Niny Khor and John Pencavel explore several ways of measuring income mobility. In terms of inclusive growth, the existence of income mobility over a longer period of time may mitigate the impacts of widening income inequality measured using cross-sectional data. The authors found considerable income mobility in the People's Republic of China, with income mobility lower among rural households than among urban households. Social welfare functions are posited that allow for a trade-off between increases in income and increases in income inequality. Results suggest strong increases in well-being for urban households while the corresponding changes in rural households are much smaller.

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