Temporal and Regional Variation in Earnings Inequality:

Urban China in Transition between 1988 and 1995*

Seth M. Hauser

Yu Xie

Forthcoming in *Social Science Research*

10/30/03

Department of Sociology and Population Studies Center

University of Michigan

WORD COUNT: 11,000

* Please address all correspondence to Seth M. Hauser (email: hausers@umich.edu) or Yu Xie (email: yuxie@umich.edu), Population Studies Center, University of Michigan, 426 Thompson St., PO Box 1248, Ann Arbor, MI 48106-1248.

Temporal and Regional Variation in Earnings Inequality: Urban China in Transition between 1988 and 1995

Abstract

This article examines trends in earnings inequality in urban China between 1988 and 1995, focusing on regional variations in earnings determination and economic growth. Specifically, we examine how returns to human capital and political capital changed during this period and how these changes varied across cities experiencing differential rates of economic growth. We find that net returns to schooling almost doubled for both men and women during this period. Returns to party membership, net of other factors, more than doubled. However, these increases in returns to human capital and political capital do not account for a sharp rise in the overall level of inequality. Consistent with earlier research based on cross-sectional data, increases in returns to schooling were depressed in cities experiencing greater levels of economic growth in the intervening period.

Temporal and Regional Variation in Earnings Inequality: Urban China in Transition between 1988 and 1995

Introduction

China's recent economic reforms, designed to stimulate economic output, provide stratification researchers the opportunity to measure the impact of economic change on social and economic inequality. Discussions of such changes under China's reforms have been dominated by market transition theory, which predicts the gradual replacement of politics by markets as the main mechanism in generating social and economic inequalities (Nee 1989, 1991, 1994, 1996, 2001). We will refer to this theory as the "marketization" hypothesis. A large body of literature in sociology has merged to test this hypothesis (see e.g. Bian and Logan 1996; Parish and Michelson 1996; Walder 1990, 1992a, 1992b, 1995, 1996; Walder *et al.* 2000; Xie and Hannum 1996, Zhao and Zhou 2002; Zhou 2000). However, because marketization itself cannot be directly observed and has not been satisfactorily operationalized (Walder 1996, 2002), stratification researchers have had difficulty determining whether their empirical results support or reject the marketization hypothesis.

We are no different. As we will report in detail in this paper, we have found unambiguous increases in the earnings returns to both education and party membership. Similar findings have been reported by Yu (1998), Zhou (2000), and Zhao and Zhou (2002). While these changes can clearly be attributed to economic reforms in general, it is not clear whether they can be attributed to the "marketization" forces hypothesized by market transition theory.

There are three approaches to operationalizing marketization: institutional, segmentation, and contextual. The institutional approach is based on the insight that, as a legacy of the socialist economy, *danwei*, or work units, continue to play a crucial role in the economic and social lives of urban residents in China (Bian 1994; Bian and Logan 1996; Tang and Parish 2000; Walder 1992; Xie and Hannum 1996).[1] The basic idea is that marketization affects workers primarily through their work units, which vary in terms of proximity to market forces (Wu 2002). One disadvantage of the institutional approach is that it requires the collection of firm-level data. To avoid this data demand, scholars have resorted to the segmentation approach, which measures the marketization effect on a worker by individual-level labor market attributes, such as sector of employment (e.g., Zhao and Zhou 2002; Zhou 2000) and type of work organization (Bian and Logan 1996; Zhou 2000). This approach assumes that workers in non-state segments of the labor market are affected more by marketization than those working in the segments owned or controlled by the state.

While the institutional and segmentation approaches are appealing, actual measures derived from them suffer from a methodological problem: they presume the exogenous sorting of workers into work units or segments. One of the primary functions of a market is the efficient allocation of resources and goods according to supply and demand. Thus, in the presence of a labor market, the sorting of workers into different work units or segments should be endogenous rather than exogenous to labor market outcomes (such as earnings). As a result, it is not possible to separate institutional—or market segment—effects from the effects of the selection

[1] There is no functional equivalent to *danwei* in the United States. Traditionally, the work unit has been the fundamental organization through which the Communist Party and the government exercised control over people's social, economic, and political lives. For example, until recently, housing was allocated through these work units, as was permission to marry and bear or adopt children.

mechanisms which sort workers into different work units or market segments (Wu and Xie 2003). The private sector of employment, for example, likely consists of workers who came from two vastly different sources: those who were "pushed," and those who "jumped."[2] At the low end of the spectrum, former employees in state enterprises may be pushed to the private sector through lay-offs, early retirements, and/or underemployment (thus underpay). They are likely to have low human capital. At the high end of the spectrum, former cadres and educated professionals may voluntarily give up their "iron rice bowl" in the state sector for very high returns in the private sector ("plunge into the sea"). The second group is very different from the first. Combining them would artificially inflate the return to education in the private sector.

Recognizing that labor markets are essentially local and that the pace of economic reforms has been regionally uneven, Xie and Hannum (1996) proposed the contextual approach: measuring marketization with the rate of economic growth within the confines of a geographic area. One potential threat to the contextual approach is internal migration. However, city-to-city migration requires physical relocation and thus a substantial opportunity cost (including the cost of obtaining the government's permission) and thus is much less likely than a change of employment sector or type of employer within the same locale. While some scholars have questioned the validity of this contextual approach because it risks conflating institutional change with economic output (Nee 2001; Szelenyi and Kostello 1996; Walder, 1996, 2000), it is unclear whether these concepts are empirically separable: a recent attempt by a group of young economists in China to develop marketization indices for different regions in China found all available indices of marketization to be highly correlated with GDP (National Economic

---

[2] We borrow these phrases from the title of a book on constraints of human rationality by Gambetta (1987), *Were They Pushed or Did They Jump?*

Research Institute 2001).[3] In this paper, we adopt Xie and Hannum's (1996) contextual approach to operationalizing marketization.

Moreover, our ultimate interest lies in the impact of economic reforms on social and economic inequality, not in the marketization hypothesis *per se*. The direct impact of macro-level economic development on individual-level economic opportunity and hence on inequality has long concerned economists (Kuznets 1955). In particular, economists have hypothesized that economic growth itself produces higher returns to schooling because "individuals who are more efficient resource allocators will be better able to take advantage of the changed opportunity sets" (Chiswick 1971, p.28). Our findings show how individual-level earnings determinants respond to economic growth, a matter of considerable theoretical interest regardless of whether it can be said to confirm or disprove the marketization thesis.

In particular, we are interested in testing Chiswick's prediction (1971) that a period of economic growth will be accompanied by an increase in the return to schooling. If this hypothesis is correct, we would expect higher increases in the return to schooling in those cities experiencing the greatest economic growth, as highly educated individuals take advantage of new opportunities in a booming economy. With an analysis of 1988 data, however, Xie and Hannum (1996) found that the city-specific return to schooling was actually lower in cities that

[3] Their work examines measures of marketization in five broad domains: relationship of government and market institutions, development of the nonstate economy, development of the consumer market, financial, labor and banking markets, and the market infrastructure and legal environment. In the second domain, for example, they examined the proportion of GDP generated by the nonstate economy, the proportion of the employed labor force in the nonstate sector, and the proportion of capitalized assets in the nonstate sector.

had experienced more economic growth. Recent research on income inequality within firms (Wu 2002) reports lower returns to education for bonuses in high-profit firms than in low-profit firms, suggesting that adherence to the socialist ethos of equality is stronger where resources are plentiful. If this hypothesis can be generalized to the city level, it predicts the opposite result from Chiswick: higher economic growth may be associated with a lower return to education or slow down a long-term trend of an increasing return to education. Because our research design allows us to distinguish baseline variation in the return to schooling in 1988 from trends therein between 1988 and 1995, we can test this proposition.

Related Empirical Literature

The literature on trends in overall economic inequality has focused mainly on income inequality across households, for two reasons. First, for researchers who are concerned with economic welfare, the examination of income at the household level is appropriate. Second, since the household is the basic economic unit in agriculture, researchers studying economic inequality in rural China have no option but to use household income; this also makes it necessary to study household income in order to compare or combine results across the urban/rural divide.

There has been a sizable body of research on trends in income inequality in China, most of which has appeared in economic literature. Although precise estimates vary, the consensus is that macro-level inequality has increased significantly following the implementation of economic reforms. With nationally representative household surveys, for example, Khan and Riskin (1998) report that the Gini coefficient increased from 0.233 to 0.332 in urban China between 1988 and 1995, and from 0.338 to 0.416 in rural China over the same period. These are significantly higher than the levels (between 0.16 and 0.19) observed for the years preceding or immediately following the introduction of the economic reform in 1978 (Adelman and Sunding

1987; Li 1986; Li *et al.* 1997; Liu 1984; Zhao 1990, 1999).

In the literature on the distribution of earnings among individuals, researchers consistently find that returns to schooling are significantly lower in China than elsewhere. Returns to schooling are understood as the multiplicative increase in earnings associated with an additional year of schooling, *ceteris paribus*.[4] That is, a 2% return indicates that we expect a person's income to be 2% higher for each additional year of schooling (compounding), all else held constant. Most estimates for returns to schooling in China range from negative (Gelb 1990; Peng 1992; Nee 1994) or negligible (Whyte and Parish 1984; Zhu 1991) to modest (*i.e.*, 1-3 percent per year) (Walder 1990; Xie and Hannum 1996; Zhao and Zhou 2002). This contrasts to the averages of a 14.4% per year return for other developing countries (Psacharopoulos 1981) and a 7.7% per year return for developed countries. Chiswick (1971) submits an economic explanation for higher returns in developing countries, suggesting that economic growth provides new opportunities best exploited by the educated. The dominance of the state-run sector, with its centrally determined wage system, is likely responsible for the exceptionally low return to schooling in China.

Returns to membership in the Communist Party of China are substantial, although the exact estimates vary across studies: 6% by Zhou (2000), 7.6% by Xie and Hannum (1996), and 9% by Walder (1990).[5] However, this party premium may not be a "causal" effect. Returns to

---

[4] Holding all other factors constant.

[5] The Communist Party in China is very different from political parties in the United States. The Communist Party is the ruling party of the People's Republic of China; the party's complicated bureaucratic structures both parallel and overlap the actual government institutions (Lieberthal 1995). Acquiring party membership is not merely a matter of sending in an annual check, it

party membership may be due to unobserved factors related to party recruitment (Xie and Hannum 1996, p.953). Indeed, there is an on-going debate, which emerged in the context of studying post-Soviet Russia, concerning the extent to which endogenous selection into the party is responsible for the observed returns to party membership (Gerber 2000, 2001; Rona-Tas and Guseva, 2001). In other words, there may be some unobserved or unobservable characteristics that increase both the likelihood of Communist Party membership and success in the labor market. In this case, Communist Party membership may serve only as a proxy for the presence of these other characteristics, instead of generating an economic premium in and of itself.[6]

Gender inequality was relatively small in pre-reform China in comparison to Western societies. Although women tended to retire from the workforce earlier than men, labor force participation of prime-aged workers was nearly universal for both men and women. The implications of economic reform for gender equity in the urban labor market have been somewhat neglected, with past research more focused on employment for married women (cf. Brinton *et al.* 1995) and gender inequities in rural areas (cf. Jacka 1997). Studies using data

---

involves an extended selection process, followed by a year long probationary period. Party membership has traditionally been a prerequisite for advancement into elite positions.

[6] It may also be the case that only some portion of the party premium is due to unobserved factors, so that party membership does provide a real advantage. Liu (2003), in his analysis of data from the 1988 Chinese Household Income Project, attempts to address this endogeneity issue. Liu finds that, contrary to Gerber's results for Russia, OLS estimates of the benefit of party membership underestimate the real effect of party membership. Allowing for the mutual dependence of party membership and earnings on unobserved factors, Liu reports that the estimated effect of party membership nearly quadruples from 10% to approximately 38%.

from the late 1980s indicate that, in urban China, women earned between 80 and 90 percent as much as men (Atinc 1997; Tang and Parish 2000), which by international standards is high. Although women have continued to make gains in educational attainment, it has been suggested that profit-maximization management styles introduced by the economic reform have marginalized women workers in the labor market and thus negatively impacted the relative gender equity in earnings (Shu and Bian 2002, 2003; Tang and Parish 2000).

The Study in Context

China is not alone in having experienced increases in economic inequality over the past two decades. Researchers focusing on the U.S. have documented and attempted to explain the rise in the level of inequality that began in the 1980s (cf. Bound and Johnson 1992; Gottschalk 1997; Levy and Murnane 1992). This body of research reveals that much of the rise in inequality in the U.S. is due to increases in within-group inequality, as technology amplifies unobserved characteristics associated with productivity, while returns to formal education have also increased (Mare 1995) in the same period. On a much greater scale and for different reasons, China has undergone a similar transformation. Economic reforms have been accompanied by political opening, technological modernization, and near miraculous levels of economic growth. Increasing demand for highly skilled workers and the influx of private and foreign enterprises, not bound by centrally-determined wage systems, are likely culprits for the increasing inequality. Both an increase in the supply of technically trained personnel and an increase in returns to workers' human capital could lead to the rise in overall inequality.

In this paper, we focus on temporal and regional variations in earnings inequality in urban China between 1988 and 1995. We do not adjudicate between the effects of politics and the effects of markets. Our concern is with the larger patterns and trajectories of inequality and earnings determination as they relate to the success of economic reforms. We do not attempt to uncover all the causes underlying earnings inequality in urban China. Nor do we claim to explain fully the rise in overall earnings inequality with changes in individual-level attributes and their effects on earnings. The fundamental questions in this study remain focused and well defined: (1) Do earnings returns to workers' characteristics change? (2) How do these changes relate to trends in aggregate inequality? and (3) How are the changes associated with the regional level of economic development?

Data and Methods

In this paper, we analyze individual-level data from the urban samples of the 1988 and 1995 Chinese Household Income Projects (henceforth CHIP88 and CHIP95). The main objective is to document trends in the importance of individual-level earnings determinants and their consequences for trends in overall inequality. Although the CHIP survey was conducted in many cities (55 in 1988 and 63 in 1995), we restricted the analysis to the 35 cities where the survey was replicated. This restriction controls for regional variation, an important feature of Chinese economic reforms (Xie and Hannum 1996), and thus maintains the comparability of labor markets over time for a trend study. The selected 35 cities are situated in 10 provinces (or their equivalent units). Although we cannot claim the cities are truly representative of urban China as a whole (as they do not constitute a random sample of cities), they are diverse in terms of geographical location, size, level of economic development, and distance to the coast. See Table A2 in the Appendix for the names and the basic characteristics of the cities.

At the core of the research design of this study is the repetition of the CHIP survey in the 35 cities in 1988 and 1995. This constitutes a significant advance over earlier trend studies based on repeated cross-sections in either single cities (*i.e.*, Bian and Logan 1996) or a multi-city cross-section with retrospectively recalled income histories (Zhou 2000). The problem with a single-city design is well known: China is regionally diverse, and it is always difficult to generalize results from one city to other cities. An additional benefit to multi-city data is that the large regional variation in China can be exploited to test theoretically relevant hypotheses, as we will explain below. Retrospective data present their own difficulties. First, as Zhou (2000, p. 1147) acknowledges, "recall errors are inevitable in retrospectively collected data, especially with respect to income." Second, the labor forces for earlier periods are necessarily truncated at old ages, because they cannot be in the sample of active workers at a later point (see Zhao and Zhou 2002). Thus, the labor forces constructed from retrospective data are not comparable in age. In addition, due to migration and urbanization, some workers observed in a later survey may have worked in other cities or even in the countryside.[7]

Sampling Design

Although the CHIP data are considered among the best large-scale China data sets, there is certainly room for improvement.[8] Nowhere is this more evident than in the sampling procedures (described in documentation available online from ICPSR). (See Griffin and Zhao (1993) and Riskin *et al.* (2000)). The urban data used in this paper are from a stratified subsample of the urban household surveys conducted by China's State Statistical Bureau (SSB). The primary

---

[7] Zhou's research design, however, has a distinct advantage of being able to control for between-person heterogeneity, as he had repeated observations of the same workers. Essentially, Zhou's research design allowed him to ask the question of how economic reforms have affected a given group of urban workers, rather than a broader question of how economic reforms have impacted urban workers in general.

[8] See Bramall (2001) for further discussion of the limitations of the CHIP surveys.

sampling units are *households*—not individuals. Unfortunately, the sampling procedures for the SSB's initial urban samples are undocumented. It is known, however, that the urban sample is limited to households where individuals held urban registrations (*hukou*), and that illiterates appear to be underrepresented in the sample. This suggests that the distribution of education within this sample is only a truncated representation of that in the urban population as a whole. The selection mechanism based on education itself does not necessarily pose a major threat to the validity of my findings. To the extent that the data fit our statistical model (to be elaborated in the next section), the underrepresentation of illiterates does not bias the estimation of returns to earnings– a situation akin to stratified sampling based on education.

That said, the underrepresentation of illiterates and the absence of urban residents without urban household registration suggests that we may be grossly underestimating the extent of urban earnings inequality in the aggregate. However, any findings in this paper will still be robust for the subset of the urban population included in the CHIP data. The limitation of the sample to residents holding urban registration is beneficial in that, given the significant obstacles to official migration in China, selective migration is unlikely to significantly alter the composition of this segment of each city's labor force between 1988 and 1995. We assume that any systematic omissions from the initial SSB samples were similar in character both across cities and waves.

Measures

For the CHIP88 and CHIP95 data, we constructed an earnings measure that is the sum of all sources of yearly-earned income, including cash bonuses/subsidies and income from family business, denoted as Y.[9] Earnings in 1995 were adjusted by the appropriate deflation factors so that all analyses are comparable in constant 1988 Yuan (China Statistics Press 2000). For convenience, we treat all members of sampled households who are between ages 20 and 59 and active in the labor force as independent observations. After excluding cases with missing data or earnings less than 100 1988 Yuan, the data yield 12,885 and 7,536 workers from the 1988 and 1995 waves respectively.

Besides earnings (Y), city (K), and year (T), the other determinants of earnings used in my analysis are years of schooling (X_1), work experience (X_2), membership in the Communist Party of China (X_4), and gender (X_5). Years of schooling (X_1) are derived from a categorical measure of levels of educational attainment (less than 3 years of schooling = 1; three years of schooling but less than primary school = 4; primary school = 6; lower middle school = 9; upper middle school = 12; trade school = 13; community/technical college = 15; and college and graduate school = 17).[10] Work experience (X_2) is approximated by the difference between the

---

[9] Salary and bonuses/subsidies were measured monthly on the survey and were converted to the yearly scale. Income measures were obtained specifically for each individual in the household. It is not clear how households divided income in their reports of income from family businesses. This component of income is salient for only a small percentage of urban households, and its exclusion from the analyses yields similar results. See Xie and Hannum (1996) for more discussion.

[10] There are four reasons for using a linear function of education in our analysis. First, human capital theory requires that education be considered in years of schooling as a primary source (and thus cost) of investment (Mincer 1974). In this framework, with log of earnings as the dependent variable, the coefficient for years of schooling can be interpreted readily as the rate of

current age and the age at first year of work, which varies with education (primary school and lower = 14; lower middle school = 16; upper middle school = 19; trade school = 20; community/technical college = 22; and college and graduate school = 24). As in earlier studies (e.g., Bian and Logan 1996; Li and Walder 2001; Xie and Hannum 1996), we interpret membership in the Communist Party of China (X_4) as a proxy for political capital (yes = 1). Gender (X_5) is a dummy variable (female = 1).

Table A1 of the Appendix presents the basic descriptive statistics for the variables, Y, X_1, X_2, and X_4, by gender and survey year. It shows rather large gender disparities in earnings as well as in all other determinants—years of schooling, work experience, and party membership. In Table A2, we provide the sample size and the mean earnings at the city-level by year. Table A2 also presents the maximum-likelihood estimates of the Gini inequality measures based on individual earnings ($Gini_{mle}$) and their changes over time ($\Delta Gini_{mle}$). The maximum-likelihood estimate of the Gini coefficient is calculated as

$$Gini_{mle} = 2\ \Phi[S_{log(y)}/(2^{1/2})] - 1$$

---

return and compared cross-nationally (Psacharopoulos 1981). Second, a one-degree-of-freedom specification for education effects allows me to conveniently analyze temporal and geographic variation in returns to schooling. Third, Xie and Hannum's (1996) study with the CHIP88 data shows that a linear specification fits the data very well. Fourth, a specification analysis comparing the fit of models where education is operationalized categorically, as opposed to linearly, shows that the deviations from linearity are relatively insubstantial in both 1988 and 1995.

where $S_{\log(y)}$ is the standard deviation of log(y), and $\Phi(\cdot)$ is the cumulative distribution function for a standard normal variable.[11] However, the reader should not give too much credence to the city-level measures of Gini coefficients and their changes, as they are based on small samples and have large sampling errors. For the whole sample pooled across the 35 cities, the estimated Gini coefficient increased from 0.252 in 1988 to 0.319 in 1995. These figures are similar in magnitude to those reported for household income for all cases in the two CHIP surveys (Khan and Riskin 1998).

Methods

Our statistical analysis has three steps. First, we estimate a modified human capital model with human capital, political capital, and gender as determinants of earnings for each period. This analysis reveals the changes in the effects of these determinants between 1988 and 1995. Second, exploring further the implications of the individual-level baseline earnings model, we decompose the increase in the overall level of inequality (as measured by Gini indexes) into three components: (a) a component that is due to changes in the distribution of the earnings determinants, (b) a component that is due to changes in the returns to these determinants, and (c) a residual component unexplained by the baseline model. We further decompose the residual component into a portion due to between-city variation and a portion due to within-city variation.

---

[11] See Xie and Hannum (1996) for an extended discussion of alternative Gini estimates. We computed both the sample analog and the maximum-likelihood estimates but chose to present the maximum-likelihood estimates. The differences between the two estimates are small and likely to be attributable to sampling errors. The last two columns of Table A2 give the "residual Gini" coefficients based on residuals from the baseline regression (equation 1) at the city-level, which are calculated by applying the equation above to the RMSE.

Finally, we explore whether the changes in the returns to the earnings determinants vary geographically and, if they do, whether a city-level indicator of economic growth can explain this variation. Economic growth (z), shown by city in Table A2, is operationalized as the logarithm of the ratio in per capita GDP for each city between 1988 and 1994.[12] As in Xie and Hannum (1996), we interpret the measure of economic growth as an indicator of the *success* of urban economic reforms between 1988 and 1995.[13] We use the growth indicator in a multi-level model of earnings.

Trends in the Determinants of Earnings

What happened to the effects of human capital and political capital between 1988 and 1995? We begin with the analysis of a modified human capital model borrowed from Xie and Hannum (1996). Omitting subscripts denoting the i^{th} person in the t^{th} period, we have

$$\log(Y) = \beta_0 + \beta_1 X_1 + \beta_2 X_2 + \beta_3 X_2^2 + \beta_4 X_4 + \beta_5 X_5 + \beta_6 X_1 X_5 + \varepsilon, \qquad (1)$$

where ε represents the residual unexplained by the baseline model. The β parameters are regression coefficients measuring the "return to," or the change in log-earnings (log(Y)) associated with, a unit change in an independent variable. Equation 1 is based on the classic human capital model of Mincer (1974) which includes education (X_1), experience (X_2), and experience-squared (X_2^2), with the addition of an indicator of political capital measured by party

[12] Thus, $z = \log(GDP_{94}/GDP_{88})$, where GDP is per capita gross domestic product. This is related to the growth rate r in that $r = (GDP_{94} - GDP_{88})/GDP_{88} = \exp(z) - 1$.

[13] As Xie and Hannum (1996, p.965) put it, "z is not an 'intention' measure but an 'outcome' measure."

membership (X_4), gender (X_5), and an interaction term to allow the return to schooling to vary by gender (X_1 X_5).[14]

For convenience, we re-express equation 1 below in vector/matrix notation:

$$\log(Y) = \beta' X + \varepsilon, \quad (2)$$

where $X' = [1\ X_1\ \ X_2\ \ X_2^2\ \ X_4\ \ X_5\ \ (X_1 X_5)]'$, and $\beta' = [\beta_0\ \beta_1\ \ \beta_2\ \ \beta_3\ \ \beta_4\ \ \beta_5\ \ \beta_6]'$. Note that in this specification, we estimate β separately for each period (year), producing period-specific estimates. This model can be expressed equivalently as:

$$\log(Y) = \beta^{*\prime} X + \delta' S + \varepsilon, \quad (3)$$

where S= tX, t is a scalar dummy variable (1995 = 1), and δ is a vector of parameters representing the interaction effects between the earnings determinants (X) and time (t).[15] We estimate the model in both specifications and present the results in Table 1.[16]

The successive columns of Table 1 present estimates of the β vector for CHIP88, the β

---

[14] The inclusion of gender is appropriate because women's labor force participation is nearly universal in China. Xie and Hannum (1996) found that the gender-education interaction follows a pattern of convergence, with women's earnings trailing substantially that of men at lower levels of education but approaching that of men at higher levels of education.

[15] Note that in equation (3), the β^* vector represents the coefficients of X for the first period (1988).

[16] These are results under the assumption of regional homogeneity and assuming the independence of individual observations within households. A multi-level model accounting for clustering by household does not change the substantive results. Results are available from the author upon request. Models treating regional homogeneity are dealt with extensively later in the paper.

vector for CHIP95, and the δ vector from pooled analysis. As expected, the results for CHIP88 are nearly identical to those from Xie and Hannum (1996), with small variations due to our inclusion of only those cities for which data were also collected in CHIP95. Consistent with expectations from human capital theory for earnings over the life course, the estimates for β_3 are negative. Thus, the experience effect is concave with positive effects on earnings early in the life course and diminishing effects toward the end of the (working) life course.[17]

The coefficient for years of schooling, β_1, indicates that men's rate of return to schooling is 2.0% per year in 1988 and nearly doubles to 3.8% per year in 1995 (exp(0.020) and exp(0.037) respectively). Inspection of δ_1, denoting change in the return to schooling between 1988 and 1995, indicates that this change is statistically significant—the rate of return to schooling among men has nearly doubled. In combination with β_1 and δ_1, coefficients β_6 and δ_6 (the gender interaction with schooling and the three-way interaction between gender, schooling, and time) reveal that women's rate of return to education increased substantially and is approximately twice that of men in both years—4.4% in 1988 and 7.7% in 1995 (exp(0.020+0.023) and exp(0.020+0.023+0.017+0.014)).[18] Thus, the estimates for 1988 are consistent with past

---

[17] By solving ($\delta \log Y / \delta X_2$) = 0 for the optimal number of years of experience, I find the optimal level of experience at 33.2 years and 29.6 years in 1988 and 1995, respectively. Given the cross-sectional nature of the data in each period, the work experience profile may reflect cohort changes. Under the assumption of no cohort changes, the estimates indicate a sharp and faster rise in earnings in 1995 than in 1988.

[18] As noted in our discussion of the data, sensitivity analyses were carried out to test the legitimacy of the linearity assumption for schooling and the schooling-gender interaction in both the 1988 and 1995 survey. These analyses indicated only limited departures from linearity in

research documenting China's relatively low rate of return to schooling. Furthermore, the estimate of change between 1988 and 1995 confirms earlier work showing an increase in the rate of return to schooling (Bian and Logan 1996; Zhao and Zhou 2002; Zhou 2000).[19]

This naturally leads us to consider the overall gender gap captured by β_5, the coefficient of gender. The presence of a significant gender-schooling interaction, β_6, requires that we not interpret β_5 in isolation. We present gender differences by education graphically in Figure 1. The four lines on the graph represent the return to schooling for males and females in both 1988 and 1995. The gender gap in earnings is greatest at lower levels of education. This presumably reflects different kinds of work available to men and women at low levels of education. The higher rate of return to schooling for women reflects the diminishing gender gap in earnings for more highly educated workers. The coefficient for change in the effect of gender, δ_5, indicates that the gender gap has widened for the least educated workers. This finding supports the expectation that economic reform may exacerbate gender inequalities in the labor market (Shu and Bian 2002, 2003; Tang and Parish 2000).

We now turn to the effect of membership in the Communist Party of China and its change over time. The estimates of β_4, the coefficient for the party membership variable, indicate that

---

each of the survey years. We chose to continue my analysis with the linear specification to facilitate the analysis of trends and regional variation, as well as maintain consistency with the human capital framework.

[19] These results implicitly assume that returns to education are the same across cohorts and, therefore, may underestimate the changing returns to education if only younger cohorts are experiencing increased returns to education. However, analysis by cohort in each period showed no significant differences between cohorts.

party members earned about 6.3% and 13.9% more than nonmembers in 1988 and 1995 respectively, net of education, experience, and gender (exp(0.061) and exp(0.130) respectively). That is, we find a dramatic increase, in fact a doubling, of the party membership earnings premium in this seven-year period. This finding is at odds with predictions of a declining significance of political capital in reform-era China. Not only does the party premium persist, the relative advantage of party members—net of education, experience, and gender—*expands* in a period of economic reform and massive economic growth. For example, whereas in 1988 a male nonmember would need to have approximately 3 additional years of schooling to acquire earnings equal to those of a party member of identical work experience and gender, in 1995 that same male nonmember would require an additional 3.5 years of schooling.[20]

Although part of the party premium may be due to selectivity of party members with unobserved productive attributes (Gerber 2000, 2001), it seems implausible that selectivity would change so rapidly in seven years to account for the doubling of the party premium. After all, the composition of the party should not have changed much in seven years. An alternative interpretation is that, while the degree of selectivity did not change much between 1988 and 1995, the returns to the selectivity increased. That is, if party membership proxies for unobserved aspects of human capital, such as ability, the apparent increase in the party premium may be due to increasing returns to ability.

Thus, under the assumption of regional homogeneity, we find considerable changes in the

[20] Females would require an additional 1.4 and 1.8 years in 1988 and 1995 respectively. These were calculated by solving for X from $\beta_1 X = \beta_4$, allowing the β's to vary by gender and year, where X is the additional number of years of schooling required by a non-party member to match the additional earnings they would expect to acquire upon joining the Communist Party.

importance of earnings determinants between 1988 and 1995. Returns to schooling increased significantly for both men and women. The gender gap in earnings expanded. The returns to party membership doubled. Thus, our findings are unambiguous; there have been both statistically and substantively significant changes in earnings determination at the individual level.

Decomposing Earnings Inequality

We now focus on the relationship between macro-level trends in overall earnings inequality and the micro-level trends in earnings determination. Conceptually, the macro-level and micro-level trends are intrinsically linked: a macro-level measure of inequality is necessarily aggregated over individual-level observations. Indeed, early conjectures about the impact of economic reforms in socialist countries were couched in terms of this linkage (e.g. Nee 1991; Parish 1984; Szelenyi and Manchin, 1987; Whyte 1986). Whereas past research has focused on either the macro-level question of whether inequality increases or decreases (cf. Khan *et al.* 1992; Khan and Riskin 1998) or the micro-level question of who gains and who loses (cf. Bian and Logan 1996; Zhou 2000), with the exception of Xie and Hannum (1996), there has been no empirical attempt to link these two lines of research. The disjuncture between the two lines of research leaves open the question of whether changes in the determinants of earnings at the individual-level (micro-level) account for the increase in inequality in the aggregate (macro-level). In the following analysis and discussion, we address this gap in the literature by explicitly examining the linkage between the micro and macro levels.

We have already documented significant changes in individual-level earnings determination and, as we noted earlier, urban China is experiencing a secular increase in the level of earnings inequality. Estimates of the Gini coefficient for the 35-city sample are 0.252 in

1988 and 0.319 in 1995. Each Gini coefficient may be interpreted as the amount of deviation from equality, which is the hypothetical scenario where total earnings were equally distributed across all workers. In the previous section, we documented nontrivial changes in returns to workers' characteristics. Our objective in this section is to explore the implications of those changes in returns at the micro-level for the trend of increasing inequality at the macro-level. That is, in attempting to answer the question of what is driving the increase in inequality, we decompose the trend in earnings inequality into several components and ascertain the extent to which each of these components, including increases in the returns to earnings determinants, has contributed to the increase in the overall level of inequality.

We partition earnings inequality into two components: structural and residual. By structural component we mean the part explained by the modified human capital model (equations 1, 2, and 3). This encompasses variations due to (a) the marginal distribution of human (and political) capital in the population and (b) the returns to those characteristics. By residual component we mean the part unexplained by the model, which includes variation due to omitted or unobserved variables and variation due to chance alone. We further decompose the residual component into (c) unexplained variation between cities and (d) unexplained variation within cities (between individuals).

From this perspective, an increase in earnings inequality can result from three sources. The first source is changes in the marginal distribution of the characteristics that determine earnings (e.g., changes in the distribution of human capital). Educational opportunities have continued to expand in China. However, it is not clear *a priori* how the expansion impacts the earnings distribution, given the educational expansion's counterbalancing effects on the variance of educational attainment (cf. Lam 1997). On the one hand, an increase in educational attainment among the younger workers creates a gradation in educational attainment by age and thus contributes positively to the variance of educational attainment. On the other hand, if the variance of education among these younger workers is smaller than the variance within older cohorts, cohort replacement works to reduce inequality in the long run (Lam and Levison 1992a, 1992b).

The second potential source of increased inequality is increasing returns to individual characteristics, e.g., increasing returns to human capital. Holding the marginal distribution of a characteristic constant, an increase in the earnings returns to that characteristic will increase inequality (Xie and Hannum, p.974). The greater the variation in the characteristic, the greater the effect of a change in the return to that characteristic on overall inequality. This mode of reasoning has a long history in the literature on comparative social mobility (e.g., Featherman, Jones, and Hauser 1975), where the research question centers on the amount of net intergenerational social mobility after differences in occupation structure between generations are accounted for. Similarly, we ask the following question: Is it that the association between workers' characteristics (such as education) and earnings that has changed, causing a change in the level of income inequality? Or, alternatively, is it that the association has remained unchanged while the underlying distribution of the characteristics in the population has changed, resulting in the increase in earnings inequality?

The third potential source of increased inequality is residual inequality. Labor economists studying the increase in U.S. income inequality during the 1980s have puzzled over the phenomenon of increasing residual inequality. In a cross-sectional, cross-national study of earnings inequality among men, Blau and Kahn (1996) find that the earnings residual accounts for nearly three-quarters of the observed difference in inequality between the United States and other countries. Residual inequality has been attributed to a number of different sources in the economics literature, including differential returns to unobserved skill and post-schooling investment in human capital (Murphy and Welch 1993). A recent examination of regional variation in within-group wage inequality in the U.S. finds that residual inequality "tends to be highest in labor markets with flexible and insecure employment conditions. Specifically, high rates of joblessness, immigration, and casualization (part-time work, temporary work, and unincorporated self-employment) exert significant positive effects on the level of residual wage inequality within labor markets" (McCall 2000, p. 426). This suggests that we may find residual inequality to be an increasingly large share of earnings inequality in urban China, as restrictions on rural-to-urban migration have relaxed, and the job security of the "iron rice bowl" has shattered. In this paper, we further exploit the multi-city data to decompose residual inequality and its trend into a between-city component and a within-city component.

This decomposition analysis is modeled after Lam and Levinson (1992a, 1992b). Applying the variance operator to equation 2, we see that variation in log(Y) can be broken down into a linear combination of the structural parameters, variances and covariances of the explanatory variables, and the variance of the error term:

$$\text{var}(\log(Y)) = \beta\Omega\beta' + \text{var}(\varepsilon), \tag{4}$$

where Y is earnings, β is a column vector of coefficients, Ω is the variance-covariance matrix of X, and ε represents the residual. We further decompose the residual variance, var(ε), into a between-city component and a within-city component:[21]

[21] Between-city variation can also be conveniently thought of as a structural component, as it is captured by dummy variables representing cities in a regression. To the extent that the between-

$$var(log(Y)) = \beta\Omega\beta' + var(\beta_{0k}) + var(\varepsilon^*), \quad (5)$$

where β_{0k} represents differences in the log of mean earnings across cities, with $var(\beta_{0k}) = \gamma^2$. We are, in effect, partialing out the variation across cities in the overall levels of income (γ^2). This allows us to isolate the within-city, or between-individual, variation in log-earnings ($var(\varepsilon^*) = \sigma^2$), net of the variation across cities. Combining these elements of the variation in log-earnings, we use the standard deviation to calculate the Gini coefficient.[22]

This decomposition analysis is based on the regression estimates for equation 3 and the sample estimate of γ^2 for the 35 cities in the sample. We present in Table 2 the decomposition of the change in earnings inequality between 1988 and 1995. Panels A, B, C, and D depict the changes in the estimated Gini coefficient due to changes in population composition, returns to characteristics, between-city residual variance and within-city residual variance respectively. Panel E shows the changes in the estimated Gini coefficient due to both residual components combined. Reading across the columns, we display the Gini coefficient computed under a hypothetical condition with only one component changing between 1988 and 1995. In Panel A, the focus is on the contribution of composition (Ω) to changes in Gini. Here we alternately fix the returns (β) and residual variances (γ^2, σ^2) at the observed values in 1988 and 1995. Similarly, Panel B presents the influence of the changes in returns (β) on Gini, with composition

---

city variation is not part of the baseline modified human capital model of equation 1, we consider it to be part of the unexplained residual in the decomposition analysis.

[22] As we stated earlier, the maximum-likelihood estimate of the Gini coefficient is calculated as $Gini_{mle} = 2\,\Phi[S_{log(y)}/(2^{1/2})] - 1$ where $S_{log(y)}$ is the standard deviation of log(y), and $\Phi(\cdot)$ is the cumulative distribution function for a standard normal variable.

(Ω) and residual variances (γ^2, σ^2) fixed. Panel C presents the influence of changing between-city variation (γ^2), with composition (Ω), returns (β), and within-city variation (σ^2) fixed. Panel D focuses on the role of within-city variation (σ^2), with composition (Ω), returns (β), and between-city variation (γ^2) fixed. Finally, the last panel shows the combined influence of the changes in both between-city (γ^2) and within-city variation (σ^2), with composition (Ω) and returns (β) fixed. Note that the main diagonal of each panel is nothing more than the observed Gini coefficients from 1988 and 1995, 0.252 and 0.319.

Panel A indicates that the changing distribution of human and political capital had a modest ameliorating effect on the level of inequality. Holding the returns (β) and residuals (γ^2, σ^2) constant at the 1988 levels, the shift in population characteristics (Ω) lowers the Gini coefficient by 0.003. For the 1995 data, if the composition distribution had remained the same as in 1988, we would have a Gini coefficient of 0.323, 0.004 higher than the Gini actually observed in 1995. That is, by itself, the change in the distribution of human and political capital in the workforce would have slightly reduced earnings inequality.

As expected, Panel B shows that changes in returns (the β vector) to schooling, experience, Communist Party membership, and gender contribute to a moderate increase in the level of earnings inequality. We observe an increase in the Gini coefficient from 0.252 to 0.269, a change of 0.017, when we keep every component at the 1988 level but allow the returns to shift to the 1995 level. Similarly, when we hold the composition and the residual variances at the 1995 level, the changes in returns between the periods increases the Gini coefficient from 0.310 to 0.319. That is, increasing returns to human and political capital exacerbated earnings inequality.

We show in Panel C that changes in between-city variation in earnings (γ^2) also contribute to an increase in the level of earnings inequality. We observe an increase in the Gini coefficient from 0.252 to 0.269, a change of 0.017, when we keep every component at the 1988 level but alter the between-city variation to the 1995 level. Similarly, holding all else constant at the 1995 level, the changes in between-city variation in earnings between the periods would increase the Gini coefficient from 0.306 to 0.319. Here we find that a portion of the increase in inequality is due to increasing differences in the distribution of earnings across cities.

Panel D reveals that the lion's share of the increase in income inequality between 1988 and 1995 is due to the rapid rise of the variance of the within-group (between-individual or within-city) residual earnings component (σ^2) unexplained by the baseline model, even after the partialing of cross-city variation. All else remaining at the 1988 levels, a shift from the 1988 to the 1995 residual variance causes an increase in the Gini coefficient of 0.046. With all other factors fixed at the 1995 levels, the increase in residual variance contributes an increase of 0.042 in the Gini coefficient from 0.277 to 0.319. The results indicate that, as is the case for the U.S., the most significant part of the increase in earnings inequality is occurring among workers with the same observed characteristics.

Panel E allows us to see the contribution of both residual components, together, to the increase in inequality between 1988 and 1995. With both the composition and returns fixed at 1988 levels, a shift in the residual variance terms causes an increase in the Gini coefficient of 0.060—from 0.252 to 0.312. This suggests that the increases in between-city and within-city variation account for nearly 90% of the actual observed increase in the Gini coefficient from 0.252 to 0.319. Of this change, approximately 75% of the combined effect of between-city and within-city variation is attributable to within-city variation.

We emphasize that the bulk of the increase in earnings inequality is not due to changes in the marginal distribution of human and political capital in the labor force, or changes in the returns to these characteristics, or changes in regional differences. If the within-city (within-group) residual variance (σ^2) and between-city residual variance (γ^2) had remained constant from 1988 to 1995, we would have observed only a modest increase in the Gini coefficient from 0.252 to 0.261 (see Panel E). Even in the absence of changes in the returns, the composition distribution, and the cross-city variation between the periods, the increase in the within-city residual variance alone would raise the Gini by 0.046 to 0.298 (Panel D). Recall that the total change in the Gini coefficient from 1988 to 1995 is an increase of 0.067 (from 0.252 to 0.319). Hence, the counterfactual exercises shown in Table 2 lead us to conclude that the dramatic increase in China's urban earnings inequality is due largely to increasing within-group variation in earnings.

What accounts for this increase in within-group variation? This is the crux of the problem. The variation in earnings among individuals with the same observed characteristics increased substantially. This may be due to increasing returns to intangibles—unobserved or unobservable characteristics of individuals—that are increasing in value in the transforming economic context. Some researchers have attributed these changes to sectoral, industry, or occupational variations, but segmented market analysis is highly problematic due to the difficulty in accounting for the processes which sort individuals into market segments (Wu and Xie 2003). Information on individuals' labor market histories, unavailable in CHIP, would be necessary to adequately address this issue. Based on the evidence here and elsewhere, the best we can do is to speculate that these trends are due to some combination of changes in the returns to unobserved characteristics, increasing importance of technology, and increasing variation across market segments.

Regional Variation in the Trends of Earnings Inequality

Application of the baseline human capital model in equation 1 by period (as presented in Table 1) provides a crude description of the general trends in earnings determination between 1988 and 1995. However, this approach ignores the reality of large regional variations in China. In the preceding decomposition analysis, we refined the baseline model by including an additive between-city component. We now turn to a more systematic analysis of the regional variation in the trends of earnings inequality.

Xie and Hannum (1996) give a detailed account of the origin and the extent of regional variation in reform-era China. To most China observers, the regional dimension is a crucial feature of the on-going economic reform, although not all researchers have appropriate data to study it. Even with data from diverse regions in China, incorporating the regional dimension into studies of earnings inequality is not straightforward. In an extreme form, if we assume that each city had its own earnings regime at the baseline period and a unique trajectory over time, we would separately estimate equation 1 by city and period. These results are presented in Appendix Table A3. This approach, however, is extremely unparsimonious and cannot explain how the earnings regime varies systematically across cities. To understand the regional variation in changes in returns to earnings determinants, we adopt a multi-level approach.

To this end, we extend the general multi-level model from Xie and Hannum (1996) to examine *trends* in earnings determinants and simultaneously account for China's vast regional heterogeneity. Following Xie and Hannum, we characterize the regional variation with an indicator of economic growth (z), measured as a monotonic transformation of the annualized growth rate in city-level per capita GDP for the period 1988 through 1994. Needless to say, Chinese cities may differ regionally in many aspects, such as natural resources, industrial bases, ties to investors/entrepreneurs in Hong Kong, Taiwan, and abroad in general, economic freedom and investment from the central government, local policies, as well as the audacity and capability of the local government to carry out economic reforms. However, measuring all of these dimensions is problematic both because they are difficult to quantify and because they are sometimes related to one another. Instead, the indicator of economic growth serves as an overall approximation of the extent to which economic reforms have been successful.[23] We use the measure of economic growth as a crude indicator to encompass regional variations in earnings determination across time. Specifically, we assume the following systematic variation at the city-level:

$$\beta_{jk} = \tau_j + \nu_{jk} \tag{6}$$

$$\delta_{jk} = \alpha_j + \lambda_j z_k + \mu_{jk} \tag{7}$$

where j indexes the j^{th} element of either β or δ vector in equation 3, and k indexes the k^{th} city. Note that ν_{jk} and μ_{jk} are city-level residual terms, assumed to follow a multivariate normal distribution. In this specification, the β parameter represents the "return" to an independent variable in 1988, and the δ parameter represents the change in the "return" between 1988 and 1995. Note that economic growth (z) enters the model only as a predictor of the δ vector.[24]

---

[23] Under the assumption that the primary purpose of economic reform is to stimulate economic growth, this interpretation is justified. Although economic growth should not be directly attributable to marketization *per se*, a recent study has found that indicators of marketization are highly correlated with GDP (National Economic Research Institute 2001).

[24] Note that z measures the *change* in GDP and should not be used to predict β (coefficients in 1988). An alternative specification would be to include a measure such as GDP in 1988 to

This formulation suggests the following interpretation of the multi-level model, using the parameters referring to party membership for illustrative purposes. Coefficient β_{4k} represents the return to party membership in the k^{th} city in 1988. It can be decomposed as the sum of the average effect across cities (τ_4) plus a city-level residual term (ν_{4k}). Thus, the city-level heterogeneity in the return to party membership at the baseline, in 1988, will be captured by the random component ν_{4k}. To assess the city-level changes in the returns to party membership, we allow not only another random component (μ_{4k}) but also a systematic component due to z, as shown in equation 7. In equation 7, α_4 refers to the average change (across cities) in the return to party membership if there is no economic growth, $\lambda_4 z_k$ refers to the amount of the change associated with economic growth, and μ_{4k} refers to the variation in change at the city-level not captured by the multi-level model. This specification of trends, *i.e.*, equation 7, differs from that of the baseline year, *i.e.*, equation 6, in that we examine the extent to which city-level variations in the changes of returns are due to the city-level variation in economic growth. That is, λ_j indicates the effect of the measure of economic growth (z) on the change in returns to the j^{th} independent variable. Thus, $\lambda_j z_k$ represents the structural portion, and μ_{jk} represents the residual portion, of city-level variation in trends.

---

account for city-level variation in the returns at baseline, or to account for city-level variation in the change in returns over time. Exploration of these alternative specifications proved unfruitful. While log-GDP in 1988 accounted for a portion of the city-level variation in the level of earnings at baseline (β_0), the substantive results, with respect to change over time, stand unchanged.

Table 3 shows goodness of fit statistics for a series of nested multi-level models.[25] Panel A begins with what we call "fixed coefficients models." The first is the baseline model of regional homogeneity, model A1. In this model, we restrict $\nu_{jk} = \lambda_j = \mu_{jk} = 0$, for all j. This is equivalent to the period-specific model presented earlier in Table 1. The second model, A2, includes a variance component for the intercept in the first period (ν_{0k}), and the third model, A3, adds a variance component for the change in the intercept between the periods (μ_{0k}) and the covariance between ν_{0k} and μ_{0k}. Likelihood-ratio tests, presented in the last two columns, show that the addition of these parameters in successive models greatly improves the goodness-of-fit over preceding models, indicating large regional variations in earnings levels as well as trends therein.

Panel B expands on model A3 by first allowing returns to earnings determinants to vary across cities in 1988 (*i.e.*, adding $\nu_{jk} \neq 0$, for j=1,…6) in model B1, and then allowing changes in returns to vary across cities (adding $\mu_{jk} \neq 0$, for j=1,…6) in model B2. In model B3, we trim those random components that do not contribute to model fit, keeping only those for the constant (ν_{0k}), education (ν_{1k}), experience (ν_{2k}), gender (ν_{5k}), and the gender-schooling interaction (ν_{6k}) at

---

[25] Results are shown from two-level models. Three-level models were estimated incorporating clustering at both the city and household levels, however this does not change the substantive story. Estimates of the multi-level models were made using the full-information maximum likelihood routine in the HLM software package. This approach, as opposed to restricted maximum likelihood estimation, allows me to compare all the nested models using the difference in deviance statistics for any nested models, where the difference is asymptotically distributed as chi-square with degrees of freedom equal to the difference in the number of parameters estimated in each model.

the baseline and wave (μ_{0k}), education (μ_{1k}), experience (μ_{2k}), and gender (μ_{5k}) for the changes across time. Finally, panel C shows models nested within model B3, where we allow economic growth (z) to predict city level variance in returns. Model C1, the full model, includes the systematic multi-level component for the intercept, education, experience, and gender (*i.e.*, $\lambda_0 \neq 0$, $\lambda_1 \neq 0$, $\lambda_2 \neq 0$, and $\lambda_5 \neq 0$). To preserve parsimony, we then removed those λ parameters that were estimated to be insignificantly different from zero and present the trimmed model as model C2.

The estimates for our preferred multi-level model, C2, are shown in Table 4.[26] City-level variation in economic growth only appears to be associated systematically with changes in the level of income (as indicated by λ_0) and the returns to schooling (as indicated by λ_1). Coefficient λ_0 is estimated to be 0.494, revealing a large, positive association of the trend in the level of earnings with economic growth in a city. With the exception of two cities (Wuhu and Gejiu), the z measures for all cities are positive. The mean of z across the 35 cities is 0.365, and this translates into an increase of about 20 percent in real earnings. The negative estimate of λ_1 means that the rise in returns to education was less in cities that had experienced more rapid economic growth between 1988 and 1995. This result is consistent with Xie and Hannum's (1996) finding from the earlier cross-sectional data of CHIP88 that higher economic growth is associated with lower returns to education. The impact of economic growth on changes in returns to education is nontrivial. For the city experiencing the highest economic growth ($z = 0.979$ in Huizhou) in these data, this influence reduced the rise in returns to education by 2.1 percent, negating the otherwise positive trend (*i.e.* $\delta_{jk} \approx 0$). That is, changes in returns to

[26] Estimates of the variances and covariances of the random components are omitted from the table.

education were inversely related to the level of economic growth: the cities experiencing the fastest economic growth showed no change in the returns to education, while those experiencing the slowest growth rates showed the largest increases in returns to education. If we hold economic growth at zero in the intervening period, the effects of schooling on earnings are 2.1% per year in 1988 and 4.4% per year in 1995 for men; the corresponding figures for women are 4.4% and 8.1%.

We do not find a systematic relationship between economic growth and the trend in returns to membership in the Communist Party of China. On average, party members earn 7.8% more in 1988 and 12.3% more in 1995 than nonmembers, net of education, work experience, and gender. As shown in Table 3, returns to party membership do not vary significantly across cities, and this conclusion pertains to the random component at the baseline ($\nu_4 \approx 0$), the random component for temporal change ($\mu_4 \approx 0$), and the systematic component for temporal change ($\lambda_4 \approx 0$). In contrast, we find that the gender differences and the gender-schooling interactions do vary by city at the baseline ($\nu_5 \neq 0$, $\nu_6 \neq 0$). There is also a random component for temporal change in overall gender differences, but not for the gender-schooling interactions ($\mu_5 \neq 0$, $\mu_6 \approx 0$). However, economic growth does not hold explanatory power for the city-level variation in these trajectories ($\lambda_5 \approx \lambda_6 \approx 0$).

In sum, our multi-level analysis shows substantial regional variation in levels of earnings, temporal changes in levels of earnings, and temporal changes in returns to education. In particular, we find that economic growth is associated positively with the trend in the level of earnings and negatively with the trend in the return to education. We explore theoretical implications of these findings in the next section.

Discussion

In the trend analysis ignoring regional variation, we found that the returns to education and party membership increased substantially (approximately doubled) between 1988 and 1995. In the multi-level analysis capitalizing on uneven rates of economic growth across cities, we found that the increase in returns to education was smaller in cities experiencing faster economic growth than in cities experiencing slower growth. These two findings may appear paradoxical. How do we reconcile them?

To begin with, let us note that the two findings pertain to two different dimensions—the first to time and the second to space. Although marketization is undoubtedly reflected in time (Walder 1996), many other aspects of Chinese society and economy, such as the government's wage policy, technology, and the situation of laid-offs from state-owned firms (*xia-gang*), also change with time (e.g., Zhou 2000; Wu and Xie 2003). Regional variation in the rate of economic growth is a more limited but more precise measure of marketization, as it captures the impact of the success of economic reform at a local level. While the first finding represents a secular trend for the whole country, the second finding modifies it, providing a more accurate description of the trend depending on local context. That is to say, the faster the economic growth, the smaller the increase in returns to education.

If the increase in returns to education over time is due to marketization, we would expect it to be positively, not negatively, associated, with the rate of economic growth across cities. The fact that the two findings appear contradictory prevents us from simply attributing the increasing trend in the education return to marketization. For example, returns to education in 1988 China were anomalously low by international standards (Xie and Hannum 1996). With the government relaxing and changing its rigid wage structure (Zhou 2000), an increase in the education return may be seen as a natural outcome of a process accompanied not only by further marketization but also by many other political and economic changes at the societal level.

Still, we are left with the question of why more marketization, as indexed by economic growth, is not associated regionally with a greater increase in returns to education. To explore this question further, we report results from a correlation analysis at the macro level across the 35 cities, shown in Table 5. As expected from the multi-level analysis, we see that economic growth (z) is positively correlated with the changes in both the mean of earnings and the mean of logged earnings (0.464 and 0.411 respectively). We also see that economic growth is negatively correlated with the change in the education return (δ_1), once again confirming a result from the multi-level analysis.

An important result in Table 5 is that, contrary to expectations, the correlations between economic growth (z) and the changes in two Gini coefficients ($\Delta\text{Gini}_{\text{mle}}$ and $\Delta\text{Gini}_{\text{rmse}}$) are weakly negative rather than positive. That is, inequality is either positively associated or not associated with economic growth. Furthermore, the correlations between changes in the level of earnings ($\Delta\bar{Y}$, $\log(\Delta\bar{Y})$) and changes in the Gini coefficients are moderately negative. These findings not only challenge the conventional wisdom that fast economic development is often accompanied by rising inequality but also inform us about the relationship between macro-level economic conditions and individual-level inequality in reform-era urban China. We have found that, in the context of China's explosive economic growth, faster economic growth not only raises average earnings but also reins in earnings inequality.[27]

[27] The emphasis here is on the impact of economic growth on *changes* in inequality. We also calculated the correlations between the measure of economic growth (z) and levels of inequality

An alert China observer may draw the implication from these results that the Chinese government's strategy of fast economic development seems to work. When a city experiences fast economic growth, all workers are better off, but the differentiation of earnings by education and unobserved characteristics is weaker, not stronger. We speculate that this is probably due to the legacy of socialistic ethos from the pre-reform economy, in which adequate and relatively egalitarian distribution of goods and services by the government was expected. Although a main thrust of the economic reform is to reduce the government's role in economic spheres, this expectation from the pre-reform era still lingers. The ability of the local government to meet the expectation, however, depends on the capacity of local economy. When a local economy expands, workers benefit not only from having higher levels of average earnings but also lower degrees of social inequality (say by education). When a local economy flounders, an increase in inequality follows as a result of relative scarcity.

Taken as a whole, our analysis provides new evidence that contradicts Nee's market transition theory, which predicts that marketization reduces returns to political capital and enhances returns to human capital. We find that returns to party membership more than doubled during this period of rapid marketization and unprecedented economic growth. Moreover, there was no regional variation whatsoever in this party premium, let alone variation due to differential economic growth. Although the increase in returns to schooling is consistent with predictions derived from market transition theory, a more nuanced analysis taking into account regional heterogeneity reveals that this measure of marketization, economic growth at the city level, is

---

(Gini coefficients in 1988 and 1995) and found them to be weakly positive but close to zero (between 0.030 and 0.108).

negatively associated with an increase in returns to education.

The result concerning returns to party membership requires more discussion. It is tempting to attribute this finding to "power persistence" (Bian and Logan 1996) or "power conversion" (Rona-Tas 1994), a thesis that the pre-reform political elite stand to profit from market transition by parlaying their political position into advantaged positions in the emerging market-based institutions. It has been found, for example, that party members have advantages in promotions and other opportunities that lead to higher pay (cf. Li and Walder 2001, Walder 1995, and Walder, Li and Treiman 2000).

However, it is also possible that party membership may serve as a proxy for otherwise unobserved aspects of human capital (Gerber 2000; Xie and Hannum 1996, p.953). According to the second interpretation, returns to party membership should increase when returns to education increase. These increases may simply reflect the fact that employers now have more freedom to set wages and reward more highly all attributes with market value, including both education and party membership.

These findings are noteworthy, however, regardless of their implications for market transition theory. Economic growth, measured by GDP, has been the fundamental objective of the economic reform in China. These results unambiguously show how individuals' economic opportunities are facilitated or constrained by macro-level economic growth. By assessing the relationship between economic growth itself and inequality, my research situates our understanding of individual-level determinants of earnings within the larger economic context of rapid economic growth.

As Bian and Logan (1996), Stark (1996), Fligstein (1996), Walder (1996), Szelenyi and Kostello (1996), Xie and Hannum (1996), and Zhou (2000) have argued, the post-socialist transition is never a unidimensional evolution that gradually transforms a socialist economy from being redistribution-oriented to being market-driven. Economic reforms in China should be understood as a complex process that simultaneously involves political, economic, and social changes in ways not necessarily approaching the political economy and social structure presumed to be typical of capitalist societies. That is, marketization is just one facet of economic reforms: as such, it neither dominates other political and economic processes nor necessarily implies a gradual path towards a Weberian ideal type of a market society.

Thus, while these results have important implications for the marketization hypothesis, we believe that the field should move beyond the narrow focus on the market transition debate and study the multiple processes, be they institutional, political, or market-driven, which operate to generate, sustain, or amplify inequality in contemporary China. Different researchers may contribute toward our understanding of contemporary Chinese society by concentrating research efforts on selected aspects of the complicated and large-scale transformation associated with China's economic reforms. When empirical results are not consistent with an overarching theory, they call for more concrete, nuanced theories that enable researchers to better explain and predict social transformations in China and other post-socialist societies.

Conclusions

In the preceding analyses, we have exploited a unique dataset with repeated cross-sectional surveys in 35 urban labor markets during the second decade of China's economic reforms. Three broad conclusions have emerged from the analyses concerning temporal and regional variations in earnings inequality in contemporary urban China. First, this study provides new and unambiguous evidence of changes in individual-level earnings determinants. In particular, returns to education and party membership have increased, and the gender gap has widened.

Second, we show that the rising tide of earnings inequality is mainly driven by increases in within-group earnings variation. That is, the unexplained variation in earnings for workers with the same observed characteristics has increased substantially. Third, concerning regional variations in trends, we find that the pace of economic growth is associated positively with earnings levels but negatively with the increase in returns to education. In contrast, changes in returns to party membership and gender differences do not vary systematically with economic growth. We do find, however, that there were substantial increases in gender inequality between 1988 and 1995.

In conclusion, we reiterate our view that contemporary China defies simple characterizations, however appealing they may appear. This study of the temporal and regional variation in earnings inequality has yielded empirical results that cannot be reconciled with predictions based on either redistributive or market economies in their ideal-typical forms. We note that the sociological literature on China has blossomed in recent years, but the literature has suffered from the over-emphasis of a single macro-institutional theoretical paradigm—the market transition theory developed by Nee (1989, 1991, 1994, 1996, 2001). The debate has served to inspire some lively discussion in the literature. The question of whether changing stratification in China can be attributed to marketization, government policy, or some other factors is central to sociological inquiry. However, this pursuit is hampered by formidable methodological obstacles and theoretical barriers. Methodologically, there is no easy way to operationalize marketization that would allow the unique identification its effects apart from those of other factors. Theoretically, there is such conceptual ambiguity and flexibility in market transition theory that the debate is unlikely to be settled by empirical research. In fact, we assert that the problem is not with market transition theory *per se* but with an underlying premise thereof: that China's future path can be predicted well in advance from sociological reasoning alone.

To understand the social consequences of economic reform in China, researchers undoubtedly need more refined and more targeted theories. At the same time, we suggest that we not let our search for a better theory overshadow our commitment to documenting and explaining the empirical social phenomena themselves. We call for sociological research around the study of social phenomena, combining empirical research and theoretical development, rather than simply confirming or rejecting a single overarching theory. In this research, we expressly set out to describe the temporal and regional variation in earnings inequality. In doing so, our study has yielded some interesting empirical results describing China's recent economic and social transformation. Returns to education and, contrary to theoretical predictions, returns to Communist Party membership have increased dramatically. These increasing returns, however, do not account for increasing earnings inequality. Rather, the trend of increasing earnings inequality is due to increases in within-group variation in earnings. Finally, we reject the supposition, derived from Chiswick (1971), that higher economic growth is associated with higher returns to schooling. Somehow the socialist ethos of equality is best maintained where economic growth is higher and resources are plentiful. We have been cautious in interpreting our empirical results but hope that others will derive new theoretical interpretations from them and supplement them with further research into this and related topics.

References

Adelman, Irma and David Sunding. 1987. "Economic Policy and Income Distribution in China." *Journal of Comparative Economics* 11(3): 444-461.

Atinc, Tamar Manuelyan. 1997. *Sharing Rising Incomes: Disparities in China.* Washington, D.C.: World Bank.

Bian, Yanjie. 1994. *Work and Inequality in Urban China.* Albany, NY: State University of New York Press.

Bian, Yanjie and Logan, John R. 1996. "Market Transition and the Persistence of Power: The Changing Stratification System in Urban China." *American Sociological Review* 61: 739-758.

Blau, Francine and Lawrence Kahn. 1996. "International Differences in Male Wage Inequality: Institutions Versus Market Forces." *Journal of Political Economy* 104: 791-837.

Bound, John and George Johnson. 1992. "Changes in the Structure of Wages in the 1980's: An Evaluation of Alternative Explanations." *American Economic Review* 82: 371-92.

Bramall, Chris. 2001. "The Quality of China's Household Income Surveys" The China Quarterly 167: 689-705

Brinton, Mary, Yean-ju Lee, and William L. Parish. 1995. "Married Women's Employment in East Asia." *American Journal of Sociology* 100: 1099-1130.

Cao, Yang and Victor Nee. 2000. "Comment: Controversies and Evidence in the Market Transition Debate" *American Journal of Sociology* 105(4): 1175-1189.

China Statistics Press. *China Statistical Yearbook 2000*. University of Michigan China Data Center. http://www.umich.edu/~iinet/chinadata/

Chiswick, Barry. 1971. "Earnings Inequality and Economic Development." *Quarterly Journal of Economics* 85:21-39.

Featherman, David L., F. Lancaster Jones, and Robert M. Hauser. 1975. "Assumptions of Social Mobility Research in the US: The Case of Occupational Status." *Social Science Research* 4:329-360.

Fligstein, Neil. 1996. "The Economic Sociology of the Transitions from Socialism." *American Journal of Sociology* 101: 1074-1081.

Gambetta, Diego. 1987. Were They Pushed or Did They Jump?: Individual Decision Mechanisms in Education. Cambridge, MA: Cambridge University Press.

Gelb, Alan. 1990. "TVP workers' incomes, incentives, and attitudes." in *China's rural industry: structure, development, and reform.* Edited by William A. Byrd and Qingsong Lin. Oxford: Oxford University Press.

Gerber, Theodore P. and Michael Hout. 1998. "More Shock than Therapy: Market Transition, Employment, and Income in Russia, 1991-19995." *American Journal of Sociology* 104(1): 1-50.

Gerber, Theodore P. 2000. "Membership Benefits or Selection Effects? Why Former Communist Party Members Do Better in Post-Soviet Russia." *Social Science Research* 29: 25-50.

Gerber, Theodore P. 2001. "The Selection Theory of Persisting Party Advantages in Russia: More Evidence and Implications (Reply to Rona-Tas and Guseva)." *Social Science Research* 30: 653-671.

Gottschalk, Peter. 1997. "Inequality, Income Growth, and Mobility: The Basic Facts." *Journal of Economic Perspectives* 11(2): 21-40.

Griffin, Keith, and Zhao Renwei. 1993. *Chinese Household Income Project, 1988* [Computer file]. New York, NY: Hunter College Academic Computing Services, producer, 1992. Ann Arbor, MI: Inter-university Consortium for Political and Social Research, distributor, 1993.

Hsiung, Bingyuang and Louis Putterman. 1989. "Pre- and post-reform income distribution in a Chinese commune: the case of Dahe township in Hebei province," *Journal of Comparative Economics* 13: 406-445.

Jacka, Tamara. 1997. *Women's Work in Rural China: Change and Continuity in an Era of Reform*. New York: Cambridge University Press.

Khan, Azizur Rahman, Keith Griffin, Carl Riskin, and Renwei Zhao. 1992. "Household income and its distribution in China." *China Quarterly* 132:1029-61.

Khan, Azizur R. and Carl Riskin. 1998. "Income and Inequality in China: Composition, Distribution and Growth of Household Income, 1988 to 1995." *China Quarterly* 154: 221-253.

Kuznets, Simon. 1955. "Economic Growth and Income Inequality." *American Economic Review* 45:1-28.

Lam, David and Deborah Levison. 1992a. "Age, Experience, and Schooling: Decomposing Earnings Inequality in the United States and Brazil." *Sociological Inquiry* 62(2): 218-45.

Lam. David and Deborah Levison. 1992b. "Declining Inequality in Schooling in Brazil and Its Effects on Inequality in Earnings." *Journal of Development Economics* 37:199-225.

Lam, David. 1997. "Demographic Variables and Income Inequality." University of Michigan Population Studies Center Research Report no. 97-385.

Levy, Frank and Richard J. Murnane. 1992. "U.S. Earnings Levels and Earnings Inequality: A Review of Recent Trends and Proposed Explanations." *Journal of Economic Literature* 30:1333-81.

Li, Bobai and Andrew Walder. 2001. "Career Advancement as Party Patronage: Sponsored Mobility into the Chinese Administrative Elite, 1949-1996." *American Journal of Sociology* 106: 1371-1408.

Li, Chengrui. 1986. "Guanyu Zhongguo jinjiniande jingji zhengce dui jumin shouru he xiaofei zhuangkuang yingxiang de tongji baogao." *Jingji Yanjiu* January No. 1.

Li, Shi, Zhao Renwei and Zhang Ping. 1997. "The Changes of China's Income Distribution During the Transition." Dissertation for the International Workshop in Beijing, Aug. 1997.

Lieberthal, Kenneth. 1995. *Governing China: from revolution through reform*. New York: W.W. Norton & Company, Inc.

Liu, Guoguang, editor. 1984. Zhongguo Jingji Fazhan Zhanlue Wenti Yanjiu (The Study of Some Questions Concerning China's Development Strategy). Shanghai.

Liu, Zhiqiang. 2003. "The Economic Impact and Determinants of Investment in Human and Political Capital in China." *Economic Development and Cultural Change*

Mare, Robert D. 1995. "Changes in Educational Attainment and School Enrollment." Pp. 155-214 in *State of the Union: America in the 1990s,* vol. 1, *Economic Trends,* edited by Reynolds Farley. New York: Russell Sage Foundation.

McCall, Leslie. 2000. "Explaining Levels of Within-Group Wage Inequality in U.S. Labor Markets." *Demography* 37(4): 415-30.

Mincer, Jacob. 1974. *Schooling, Experience and Earnings*. New York: Columbia University Press.

Murphy, Kevin and Finis Welch. 1993. "Industrial Change and the Rising Importance of Skill," in S. Danziger and P. Gottschalk, eds., *Uneven Tides: Rising Inequality in the 1980's*, New York: Russell Sage Foundation, pp. 101-132.

National Economic Research Institute. 2001. *Zhonguo Shichanghua Zhishu (China's Marketization Index)*. China: Economic Sciences Publishers.

Nee, Victor. 1989. "A Theory of Market Transition: From Redistribution to Markets in State Socialism." *American Sociological Review* 54:663-681.

Nee, Victor. 1991. "Social Inequalities in Reforming State Socialism: Between Redistribution and Markets in China." *American Sociological Review* 56:267-282.

Nee, Victor. 1994. "Markets and Inequality: Why Marx and Smith are Both Right." Working Paper no. 94.4. Mario Einaudi Center for International Studies, Cornell University.

Nee, Victor. 1996. "The Emergence of a Market Society: Changing Mechanisms of Stratification in China." *American Journal of Sociology* 101:908-49.

Nee, Victor. 2001. "Post-Socialist Stratification" in *Social Stratification: Class, Race, and Gender in Sociological Perspective*, 2nd Edition, edited by David B. Grusky. Boulder: Westview Press.

Parish, William L. 1984. "Destratification in China." in *Class and Social Stratification in Post-Revolution China.* Edited by J. Watson. New York: Cambridge University Press.

Parish, William L. and Ethan Michelson. 1996. "Politics and Markets: Dual Transformations." *American Journal of Sociology* 101:1042-1059.

Peng, Yusheng. 1992. "Wage Determination in Rural and Urban China: A Comparison of Public and Private Industrial Sectors." *American Sociological Review* 57(2): 198-213.

Perkins, Dwight H. 1986 "The Prospects for Chinas Economic Reforms" in *Modernizing China: Post-Mao Reform and Development*. Edited by A. Doak Barnett and Ralph N. Clough. Boulder, CO, Westview Press.

Perkins, Dwight H. 1988. "Reforming China's economic system." *Journal of Economic Literature* 26:601-645.

Perkins, Dwight H. 1994. "Completing China's move to the market." *Journal of Economic Perspectives* 8(2): 23-46.

Psacharopoulos, George. 1981. "Returns to Education: An Updated International Comparison." *Comparative Education* 17:321-41.

Riskin, Carl. 1987. China's political economy—the quest for development since 1949. New York: Oxford University Press.

Riskin, Carl, Zhao Renwei, and Li Shi. 2000. *Chinese Household Income Project, 1995* [Computer file]. ICPSR version. Amherst, MA: University of Massachusetts, Political Economy Research Institute, producer, 2000. Ann Arbor, MI: Inter-university Consortium for Political and Social Research, distributor, 2000.

Robinson, W.S. 1950. "Ecological Correlations and the Behavior of Individuals." *American Sociological Review* 15(3): 351-357.

Rona-Tas, Akos. 1994. "The First Shall Be Last? Entrepreneurship and Communist Cadre in the Transition From Socialism." *American Journal of Sociology* 100:40-69.

Rona-Tas, Akos and Alya Guseva. 2001. "The Privileges of Past Communist Party Membership in Russia and Endogenous Switching Regressions: A Reply to Gerber." *Social Science Research* 30: 641-652.

Selden, Mark. 1993. *The political economy of Chinese development.* Armonk, N.Y.: M. E. Sharpe.

Shu, Xiaoling and Yanjie Bian. 2002. "Intercity Variation in Gender Inequalities in China: Analysis of A 1995 National Survey." *Research in Social Stratification and Mobility* 19: 267-307.

Shu, Xiaoling and Yanjie Bian. 2003. "Market Transition and Gender Gap in Earnings in Urban China." *Social Forces*.

Stark, David. 1996. "Recombinant Property in East European Capitalism." *American Journal of Sociology*. 101:993-1027.

Szelenyi, Iván and Manchin, Robert. 1987. "Social Policy under State Socialism." in *Stagnation and Renewal in Social Policy.* Edited by G.E. Anderson, L. Rainwater, and M. Rein. White Plains, NY, Sharpe.

Szelenyi, Iván and Eric Kostello. 1996. "The Market Transition Debate: Toward a Synthesis?" *American Journal of Sociology* 101:1082-1096.

Tang, Wenfang and William L. Parish. 2000. *Chinese Urban Life Under Reform: The Changing Social Contract.* New York: Cambridge University Press.

Trescott, Paul B. 1985. "Incentives Versus Equality: What Does China's Recent Experience Show? " *World Development* 13(12): 205-217.

Walder, Andrew G. 1987. "Wage reform and the web of factory interests." *China Quarterly* 109:22-41.

Walder, Andrew G. 1990. "Economic Reform and Income Distribution in Tianjin, 1976-1986." in *Chinese Society on the Eve of Tiananmen.* Edited by Deborah Davis and Ezra F. Vogel. Pp. 135-56. Cambridge, MA: Harvard University Press.

Walder, Andrew G. 1992a. "Property rights and stratification in socialist redistributive economies." *American Sociological Review* 57:524-539.

Walder, Andrew G. 1992b. "Urban Industrial Workers: Some Observations on the 1980s." in *State and Society in China: The Consequences of Reform*. Edited by Arthur G. Rosenbaum. Pp. 103-120. Boulder, CO: Westview Press.

Walder, Andrew G. 1995. "Career Mobility and the Communist Political Order." *American Sociological Review* 60:309-328.

Walder, Andrew G. 1996. "Markets and Inequality in Transitional Economies: Toward Testable Theories." *American Journal of Sociology*. 101:1060-1073.

Walder, Andrew G. 2002. "Markets and income inequality in rural China: Political advantage in an expanding economy." *American Sociological Review* 67:231-253.

Walder, Andrew G, Bobai Li, and Donald J Treiman. 2000. "Politics and Life Chances in a State Socialist Regime: Dual Career Paths into the Urban Chinese Elite, 1949 to 1996." *American Sociological Review* 65:191-209.

Whyte, Martin King and William L. Parish. 1984. *Urban Life in Contemporary China.* Chicago: University of Chicago.

Whyte, Martin K. 1986. "Social Trends in China: The Triumph of Inequality." in *Modernizing China: Post-Mao Reform and Development*. Edited by A. Doak Barnett and Ralph N. Clough. Boulder, CO, Westview Press.

Wu, Xiaogang. 2002. "Work Units and Income Inequality: The Effect of Market Transition in Urban China." *Social Forces* 80(3): 1069-1099.

Wu, Xiaogang and Yu Xie. 2003. "Does the Market Pay Off? Earnings Returns to Education in Urban China." *American Sociological Review* 68: 425-442.

Xie, Yu and Emily Hannum. 1996. "Regional Variation in Earnings Inequality in Reform-Era Urban China." *American Journal of Sociology* 101:950-992.

Xu, Lizin Colin and Heng-fu Zou. 2000. "Explaining the changes of income distribution in China." *China Economic Review* 11:149-170.

Yao, Shujie. 1999. "Economic growth, income inequality and poverty in China under economic reforms." *The Journal of Development Studies* 35(6): 104-130.

Yu, Xuejun. 1998. "Earnings and Returns to Human Capital Investments During Economic Transition in Urban China." Unpublished manuscript.

Zhao, Renwei. 1990. "Income Distribution." In *The Chinese Economy and its Future: Achievements and Problems of Post-Mao Reform.* Edited by Peter Nolan and Fureng Dong. Pp. 187-199. Cambridge, MA, Polity Press.

Zhao, Renwei. 1999. "Increasing Income Inequality and its Causes in China." University of Oxford Applied Discussion Paper Series No. 213.

Zhao, Wei and Xueguang Zhou. 2002. "Institutional Transformation and Returns to Education in Urban China: An Empirical Assessment." *Research in Social Stratification and Mobility* 19:337-373.

Zhou, Xueguang. 2000. "Economic Transformation and Income Inequality in Urban China: Evidence From Panel Data." *American Journal of Sociology* 105(4): 1135-1174.

Zhu, Ling. 1991. *Rural Reform and Peasant Income in China: the Impact of China's Post-Mao Rural Reforms in Selected Regions*. Basingstoke: Macmillan.

Table 1: Models for Log-Earnings in 1988 and 1995 Assuming Regional Homogeneity

	1988			1995			(1995 vs. 1988)		
Independent Variable	β	SE		β	SE		δ	SE	
Intercept	6.753	0.022	***	6.882	0.044	***	0.129	0.045	**
Years of schooling (X_1)	0.020	0.002	***	0.037	0.003	***	0.017	0.003	***
Experience (X_2)	0.044	0.001	***	0.050	0.002	***	0.006	0.002	*
Experience^2 (X_2^2)	$(-6.63)10^{-04}$	$(2.97)10^{-05}$	***	$(-8.45)10^{-04}$	$(5.52)10^{-05}$	***	$(-1.82)10^{-04}$	$(5.77)10^{-05}$	**
Party member (1 = yes) (X_4)	0.061	0.009	***	0.130	0.015	***	0.068	0.016	***
Gender(1 = female) (X_5)	-0.365	0.025	***	-0.575	0.051	***	-0.210	0.052	***
Gender x years of schooling (X_1X_5)	0.023	0.002	***	0.037	0.004	***	0.014	0.004	***
Root mean square error	0.384			0.513					
df	12878			7529					
R^2 (%)	24.8			19.7					

Note—N = 12885 and N = 7536 for 1988 and 1995 respectively. The dependent variable is the natural logarithm of total annual earnings (1988 yuan). *** $p < .001$ ** $p < .01$ * $p < .05$

Table 2: Decomposition of Change in Overall Inequality

A. Due to Composition (Ω)	Ω_{88}	Ω_{95}	ΔGini
using β_{88}, σ^2_{88}, γ^2_{88}	0.252	0.249	-0.003
using β_{95}, σ^2_{95}, γ^2_{95}	0.323	0.319	-0.004
B. Due to Returns (β)	β_{88}	β_{95}	ΔGini
using Ω_{88}, σ^2_{88}, γ^2_{88}	0.252	0.266	+0.014
using Ω_{95}, σ^2_{95}, γ^2_{95}	0.310	0.319	+0.009
C. Due to Between-City Variation (γ^2)	γ^2_{88}	γ^2_{95}	ΔGini
using β_{88}, Ω_{88}, σ^2_{88}	0.252	0.269	+0.017
using β_{95}, Ω_{95}, σ^2_{95}	0.306	0.319	+0.013
D. Due to Within-City Variation (σ^2)	σ^2_{88}	σ^2_{95}	ΔGini
using β_{88}, Ω_{88}, γ^2_{88}	0.252	0.298	+0.046
using β_{95}, Ω_{95}, γ^2_{95}	0.277	0.319	+0.042
E. Due to Between and Within-City Variation (γ^2,σ^2)	γ^2_{88}, σ^2_{88}	γ^2_{95}, σ^2_{95}	ΔGini
using β_{88}, Ω_{88}	0.252	0.312	+0.060
using β_{95}, Ω_{95}	0.261	0.319	+0.058

Note—Entries in the table are the Gini coefficients calculated based on the following decomposition of the variance of log-earnings: $var(log(Y)) = \beta\Omega\beta^T + var(\beta_{0k}) + var(\varepsilon)$, where Ω is the variance-covariance matrix of the independent variables, and Y, β, ε are vectors of earnings, coefficients, and residuals, $var(\beta_{0k}) = \gamma^2$ is between-city variation in the level of earnings, and $var(\varepsilon) = \sigma^2$ is within-city variation in the level of earnings.

Table 3: Goodness of Fit Statistics for Multi-Level Models

Model Specification	Number of Parameters	Deviance	χ^2		*df*	Reference Model
A. Fixed Coefficients Models						
A1. Regional Homogeneity	14	24,028.25				
($\nu_{jk} = \lambda_j = \mu_{jk}=0$, for all j)						
A2. City-level Variance Component	15	20,077.59	3,950.66	***	1	A1
(add $\nu_{0k} \neq 0$)						
A3. City-level by Time Variance Component	17	19,412.89	664.70	***	2	A2
(add $\mu_{0k} \neq 0$)						
B. Random Coefficients Models						
B1. City-level Random Coefficients	50	19,193.83	219.06	***	33	A3
(add $\nu_{jk} \neq 0$, for j=1,…6)						
B2. City-level by Time Random Coefficients	119	19,121.18	72.65		69	B1
(add $\mu_{jk} \neq 0$, for j=1,…6)						
B3. Trimmed Random Coefficients	59	19,146.69	266.20	***	42	A3
(restrict $\nu_{3k} = \nu_{4k} = \mu_{3k} = \mu_{4k} = \mu_{6k} = 0$)						
C. Multi-Level Models						
C1. Full, z predicting changes	63	19,137.83	8.86		4	B3
(add $\lambda_0 \neq 0$, $\lambda_1 \neq 0$, $\lambda_2 \neq 0$, and $\lambda_5 \neq 0$)						
C2. Trimmed, z predicting changes	61	19,138.29	8.40	*	2	B3
(restrict $\lambda_2 = \lambda_5 = 0$)						

*** $p < .001$ ** $p < .01$ * $p < .05$.

Note—χ2 statistic (with its degrees of freedom reported in column labeled df) tests the statistical significance of the current model versus the reference model in goodness-of-fit. Number of parameters includes the estimated variances and covariances of city-level random components.

Table 4: Estimated Parameters of the Preferred Multi-Level Model of Earnings

	Parameter		SE
Baseline coefficients:			
Intercept (τ_0)	6.762	***	0.052
Years of schooling (τ_1)	0.021	***	0.002
Experience (τ_2)	0.042	***	0.002
Experience^2 (τ_3)	$(-6.23)10^{-04}$	***	$(4.50)10^{-05}$
Party member (1 = yes) (τ_4)	0.075	***	0.008
Gender (1 = female) (τ_5)	-0.360	***	0.043
Schooling x Gender (τ_6)	0.022	***	0.003
Trend coefficients			
Wave (1 = 1995) (α_0)	-0.017		0.070
Schooling x Wave (α_1)	0.022	***	0.004
Experience x Wave (α_2)	0.005		0.003
Experience^2 x Wave (α_3)	$(-1.93)10^{-04}$	*	$(8.90)10^{-05}$
Party x Wave (α_4)	0.041	**	0.016
Gender x Wave (α_5)	-0.203	**	0.060
Schooling x Gender x Wave (α_6)	0.013	**	0.004
Micro-macro interactive coefficients			
Wave (1 = 1995) (λ_0)	0.494	***	0.125
Schooling x Wave (λ_1)	-0.021	*	0.010
Microlevel variance component:			
var(ε)	0.146		
Deviance	19138		
Df	62		

Note—Macrolevel variance components (ν's and μ's) have been omitted from the table.
*** $p < .001$ ** $p < .01$ * $p < .05$

Table 5: Correlations among Contextual Variables and Estimated Statistics across 35 Cities in China

	GDP_{88}	z	$\Delta\bar{Y}$	$\Delta\log(\bar{Y})$	δ_1	δ_2	$\Delta Gini_{mle}$	$\Delta Gini_{rmse}$
Gross Domestic Product in 1988 (GDP_{88})	1.000	0.177	0.264	0.275	0.416	0.104	0.344	0.291
Economic growth (z)		1.000	0.464	0.411	-0.196	0.228	-0.022	-0.053
Change in mean of earnings ($\Delta\bar{Y}$)			1.000	0.907	-0.017	-0.023	-0.285	-0.255
Change in logged mean of earnings ($\Delta\log(\bar{Y})$)				1.000	0.042	-0.109	-0.113	-0.034
Change in return to schooling (δ_1)					1.000	0.062	0.260	0.151
Change in return to experience (δ_2)						1.000	0.410	0.266
Change in Gini, ML estimate ($\Delta Gini_{mle}$)							1.000	0.915
Change in residual Gini ($\Delta Gini_{rmse}$)								1.000

Note—ML = maximum-likelihood estimation. Regression coefficients (δ's) refer to the least-squares estimates for each of the 35 cities as reported in table A3. Gini coefficients are from table A2.

Table A1: Descriptive Statistics by Year and Gender

	1988		1995	
	Male	Female	Male	Female
Earnings (Y, in 1988 yuan)	2,101	1,756	3,239	2,722
	(1,617)	(1,507)	(1,791)	(1,569)
Years of Schooling (X_1)	11.12	10.38	12.22	11.63
	(3.14)	(2.88)	(2.95)	(2.78)
Work Experience (X_2, in years)	20.26	18.12	20.57	18.63
	(10.69)	(9.43)	(10.16)	(9.01)
Communist Party Membership (X_4)	34.6%	12.3%	32.2%	15.9%
N	6723	6162	3998	3538

Note—Standard deviations are in parentheses below the means.

Table 1.A2: Summary Measures of Earnings Inequality by Province and City: 1988 and 1995 CHIP Micro Data

	N		Annual Earnings (1988 Yuan)		$Gini_{mle}$		$\Delta Gini_{mle}$	$Gini_{rmse}$		$\Delta Gini_{rmse}$	z
	1988	1995	1988	1995	1988	1995		1988	1995		
Total	12885	7536	1,936 (1575)	2996 (1710)	0.222	0.278	0.056	0.184	0.246	.062	
Beijing:											
Beijing	818	830	1,957 (796)	3,853 (1723)	0.220	0.270	0.050	0.187	0.252	0.065	0.461
Shanxi:											
Taiyuan	559	294	1,693 (1867)	2,403 (1142)	0.256	0.324	0.068	0.211	0.258	0.047	0.176
Datong	423	167	1,581 (990)	2,543 (1170)	0.266	0.315	0.049	0.215	0.274	0.059	0.266
Changzhi	200	178	1,584 (581)	1,825 (887)	0.237	0.334	0.097	0.195	0.287	0.092	0.319
Yangquan	207	201	1,759 (2470)	2,440 (1133)	0.260	0.285	0.025	0.225	0.246	0.021	0.415
Liaoning:											
Shenyang	619	613	1,810 (638)	2,607 (1083)	0.171	0.237	0.066	0.140	0.216	0.076	0.474
Dalian	613	301	1,894 (588)	2,985 (1361)	0.159	0.280	0.121	0.129	0.247	0.118	0.550
Jinzhou	371	148	1,650 (706)	2,304 (1219)	0.184	0.324	0.140	0.140	0.286	0.146	0.274
Jiangsu:											
Nanjing	524	325	1,732 (526)	3,054 (1402)	0.175	0.288	0.113	0.129	0.272	0.143	0.552
Wuxi	444	132	1,889 (704)	3,479 (1387)	0.184	0.246	0.062	0.148	0.215	0.067	0.497
Xuzhou	340	142	1,702 (613)	3,286 (1469)	0.193	0.227	0.035	0.147	0.216	0.069	0.751
Changzhou	176	158	1,878 (723)	2,786 (1207)	0.200	0.241	0.041	0.170	0.202	0.032	0.327
Nantong	194	185	1,926 (632)	3,123 (1224)	0.180	0.221	0.041	0.149	0.180	0.031	0.148

Table 1.A2: Summary Measures of Earnings Inequality by Province and City: 1988 and 1995 CHIP Micro Data (Continued)

	N		Annual Earnings (1988 Yuan)		$Gini_{mle}$		$\Delta Gini_{mle}$	$Gini_{rmse}$		$\Delta Gini_{rmse}$	z
	1988	1995	1988	1995	1988	1995		1988	1995		
Anhui:											
Hefei	409	167	1,785	2,662	0.220	0.283	0.063	0.180	0.247	0.067	0.183
			(982)	(1360)							
Huainan	377	151	1,717	2,002	0.282	0.332	0.049	0.229	0.302	0.072	0.347
			(1173)	(977)							
Wuhu	173	156	1,862	2,203	0.227	0.307	0.080	0.206	0.256	0.051	-0.123
			(2150)	(1109)							
Bengbu	186	165	1,920	2,178	0.278	0.346	0.067	0.230	0.318	0.088	0.057
			(2882)	(1017)							
Henan:											
Zhengzhou	384	124	1,613	2,735	0.196	0.275	0.080	0.159	0.233	0.074	0.532
			(638)	(1224)							
Kaifeng	195	167	1,439	1,964	0.208	0.270	0.062	0.188	0.229	0.041	0.007
			(665)	(917)							
Pingdingshan	182	197	1,512	2,366	0.211	0.319	0.108	0.175	0.260	0.085	0.470
			(590)	(1068)							
Xinxiang	269	140	1,518	2,238	0.231	0.301	0.070	0.192	0.274	0.082	0.234
			(584)	(946)							
Hubei:											
Wuhan	653	443	1,760	2,849	0.191	0.289	0.098	0.151	0.254	0.102	0.368
			(618)	(1272)							
Huangshi	293	178	1,756	3,141	0.214	0.235	0.020	0.179	0.212	0.034	0.104
			(1149)	(1356)							
Guangdong:											
Guangzhou	557	347	2,652	5,554	0.273	0.326	0.053	0.258	0.297	0.040	0.547
			(2513)	(2993)							
Foshan	309	93	3,539	5,899	0.255	0.237	-0.019	0.243	0.214	-0.029	0.502
			(5014)	(3195)							
Zhanjiang	197	92	2,419	4,501	0.316	0.257	-0.058	0.287	0.237	-0.049	0.441
			(3287)	(1970)							
Shenzhen	206	85	4,065	3,972	0.228	0.317	0.089	0.190	0.246	0.057	0.525
			(1540)	(2204)							
Huizhou	220	89	2,511	3,173	0.280	0.275	-0.004	0.250	0.228	-0.022	0.979
			(2692)	(1528)							
Zhaoqing	208	90	2,337	4,845	0.300	0.332	0.032	0.264	0.300	0.036	0.863
			(1887)	(2532)							

Table 1.A2: Summary Measures of Earnings Inequality by Province and City: 1988 and 1995 CHIP Micro Data (Continued)

	N		Annual Earnings (1988 Yuan)		Gini_{mle}		$\Delta\text{Gini}_{\text{mle}}$	$\text{Gini}_{\text{rmse}}$		$\Delta\text{Gini}_{\text{rmse}}$	z
	1988	1995	1988	1995	1988	1995		1988	1995		
Yunnan:											
Kunming	575	188	2,024	2,695	0.191	0.221	0.030	0.162	0.188	0.026	0.547
			(693)	(929)							
Gejiu	275	149	1,867	2,408	0.213	0.191	-0.022	0.163	0.178	0.014	-0.003
			(926)	(783)							
Dali	288	189	1,819	3,020	0.181	0.287	0.105	0.155	0.258	0.103	0.487
			(773)	(1146)							
Dongchuan	151	169	1,884	2,532	0.220	0.228	0.008	0.173	0.204	0.032	0.104
			(1308)	(906)							
Baoshan	193	173	1,762	2,564	0.171	0.211	0.040	0.135	0.200	0.065	0.298
			(818)	(1462)							
Gansu:											
Lanzhou	1,097	310	1,887	2,231	0.277	0.311	0.034	0.218	0.266	0.047	0.083
			(1456)	(977)							

Note.—N denotes the sample size, the standard deviation of earnings is in parentheses below the mean, and the Gini coefficients are derived from maximum likelihood estimates.

Table 1.A3: Estimates of Baseline Human Capital Model by Province and City

	β_0	δ_0	β_1	δ_1	β_2	δ_2	β_3	δ_3	β_4	δ_4	β_5	δ_5	β_6	δ_6
Beijing:														
Beijing	6.847	0.558	0.016	0.012	0.044	-0.005	-6.646E-04	-1.174E-04	0.033	0.117	-0.177	-0.044	0.005	0.002
	(0.090)	(0.140)	(0.006)	(0.010)	(0.005)	(0.007)	(1.120E-04)	(1.774E-04)	(0.037)	(0.049)	(0.114)	(0.177)	(0.010)	(0.014)
Shanxi:														
Taiyuan	6.452	0.162	0.028	0.012	0.042	0.036	-6.237E-04	-9.657E-04	0.143	-0.180	-0.439	-0.925	0.031	0.054
	(0.106)	(0.198)	(0.008)	(0.014)	(0.006)	(0.010)	(1.477E-04)	(2.419E-04)	(0.043)	(0.077)	(0.135)	(0.249)	(0.011)	(0.020)
Datong	6.633	0.611	0.017	-0.009	0.049	-0.019	-7.262E-04	4.729E-04	0.023	0.111	-0.924	-0.300	0.060	0.021
	(0.141)	(0.296)	(0.011)	(0.021)	(0.007)	(0.014)	(1.672E-04)	(3.239E-04)	(0.059)	(0.113)	(0.157)	(0.320)	(0.015)	(0.028)
Changzhi	6.515	-0.511	0.014	0.050	0.056	0.008	-7.654E-04	-3.707E-04	0.038	0.068	-0.729	0.199	0.059	-0.032
	(0.227)	(0.344)	(0.017)	(0.025)	(0.011)	(0.017)	(2.547E-04)	(3.974E-04)	(0.104)	(0.139)	(0.290)	(0.434)	(0.028)	(0.039)
Yangquan	6.382	0.728	0.034	-0.007	0.063	-0.041	-1.066E-03	1.072E-03	-0.091	0.123	-0.495	-0.630	0.030	0.050
	(0.191)	(0.305)	(0.016)	(0.023)	(0.011)	(0.017)	(2.765E-04)	(4.329E-04)	(0.094)	(0.137)	(0.240)	(0.367)	(0.022)	(0.032)
Liaoning:														
Shenyang	6.788	-0.051	0.021	0.025	0.034	0.008	-4.501E-04	-2.545E-04	0.076	-0.027	-0.218	0.007	0.015	-0.003
	(0.095)	(0.144)	(0.007)	(0.010)	(0.005)	(0.008)	(1.344E-04)	(2.123E-04)	(0.031)	(0.048)	(0.112)	(0.165)	(0.010)	(0.014)
Dalian	6.891	0.289	0.016	0.012	0.034	-0.002	-4.376E-04	7.090E-05	0.057	0.088	-0.168	-0.466	0.011	0.027
	(0.085)	(0.151)	(0.006)	(0.010)	(0.005)	(0.008)	(1.296E-04)	(2.016E-04)	(0.032)	(0.059)	(0.103)	(0.188)	(0.009)	(0.016)
Jinzhou	6.653	0.665	0.021	-0.031	0.038	0.010	-5.046E-04	-5.278E-04	0.098	0.245	-0.354	-0.863	0.023	0.060
	(0.129)	(0.260)	(0.010)	(0.019)	(0.007)	(0.013)	(1.719E-04)	(3.042E-04)	(0.049)	(0.091)	(0.147)	(0.278)	(0.013)	(0.025)
Jiangsu:														
Nanjing	6.739	0.371	0.014	0.024	0.046	-0.006	-7.148E-04	-1.484E-04	0.093	0.044	-0.182	-0.045	0.009	-0.002
	(0.105)	(0.192)	(0.008)	(0.013)	(0.006)	(0.009)	(1.374E-04)	(2.330E-04)	(0.046)	(0.068)	(0.137)	(0.222)	(0.012)	(0.019)
Wuxi	6.773	0.065	0.025	0.023	0.033	0.024	-3.613E-04	-6.397E-04	0.057	0.070	-0.101	-0.162	0.006	0.012
	(0.092)	(0.211)	(0.007)	(0.014)	(0.005)	(0.013)	(1.231E-04)	(3.339E-04)	(0.043)	(0.084)	(0.106)	(0.250)	(0.010)	(0.021)
Xuzhou	6.842	0.395	0.010	0.019	0.047	-0.016	-7.835E-04	3.750E-04	0.094	-0.041	-0.618	0.700	0.042	-0.052
	(0.111)	(0.232)	(0.008)	(0.016)	(0.006)	(0.013)	(1.388E-04)	(3.096E-04)	(0.052)	(0.088)	(0.125)	(0.266)	(0.012)	(0.023)
Changzhou	6.996	-0.035	0.001	0.035	0.045	0.006	-7.288E-04	-2.897E-04	0.052	0.244	-0.393	0.002	0.025	0.002
	(0.163)	(0.246)	(0.012)	(0.018)	(0.009)	(0.013)	(2.100E-04)	(3.054E-04)	(0.078)	(0.111)	(0.192)	(0.293)	(0.018)	(0.027)
Nantong	7.128	-0.009	0.002	0.033	0.031	0.018	-4.273E-04	-4.457E-04	0.110	-0.102	-0.341	-0.430	0.022	0.033
	(0.140)	(0.212)	(0.010)	(0.015)	(0.007)	(0.011)	(1.852E-04)	(2.785E-04)	(0.053)	(0.081)	(0.152)	(0.264)	(0.014)	(0.023)

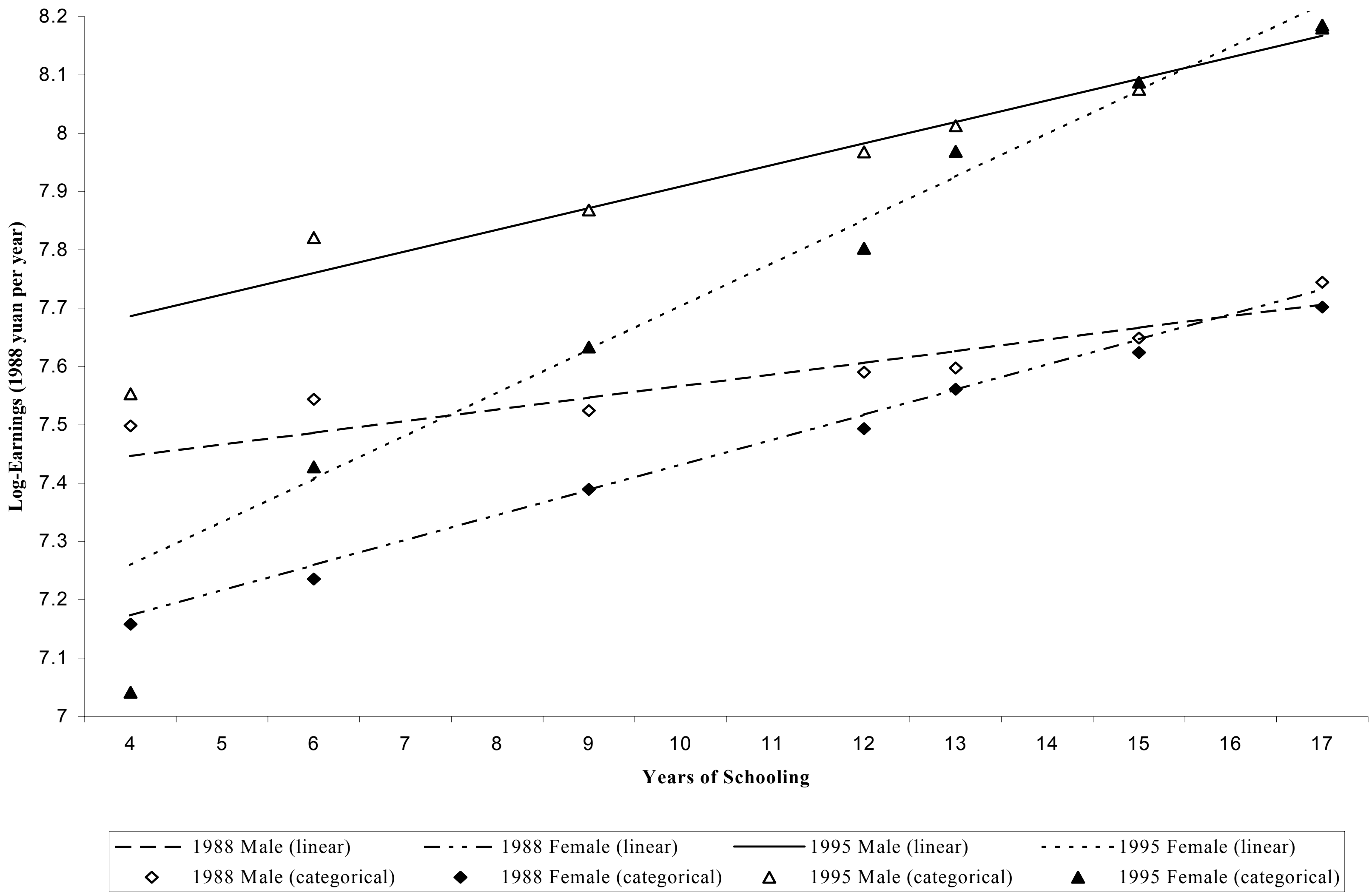
Table 1.A3: Estimates of Baseline Human Capital Model by Province and City (Continued)

	β_0	δ_0	β_1	δ_1	β_2	δ_2	β_3	δ_3	β_4	δ_4	β_5	δ_5	β_6	δ_6
Anhui:														
Hefei	6.551	0.009	0.033	0.022	0.038	0.016	-4.926E-04	-4.420E-04	0.073	-0.039	-0.155	-0.202	0.009	0.008
	(0.116)	(0.247)	(0.008)	(0.017)	(0.007)	(0.013)	(1.724E-04)	(3.411E-04)	(0.049)	(0.090)	(0.139)	(0.277)	(0.012)	(0.022)
Huainan	6.921	-0.236	0.011	0.007	0.039	0.029	-5.683E-04	-8.073E-04	0.065	0.257	-1.075	0.193	0.074	-0.011
	(0.143)	(0.332)	(0.011)	(0.024)	(0.008)	(0.018)	(2.046E-04)	(4.149E-04)	(0.068)	(0.135)	(0.157)	(0.359)	(0.015)	(0.034)
Wuhu	6.596	-0.246	0.029	0.029	0.040	0.006	-5.512E-04	-3.980E-05	0.006	0.075	-0.024	-0.501	-0.004	0.034
	(0.203)	(0.317)	(0.014)	(0.022)	(0.012)	(0.017)	(2.819E-04)	(4.012E-04)	(0.090)	(0.124)	(0.195)	(0.333)	(0.018)	(0.029)
Bengbu	6.442	0.315	0.032	-0.022	0.059	0.027	-8.788E-04	-1.056E-03	-0.022	0.129	-0.450	-0.192	0.024	0.004
	(0.249)	(0.391)	(0.020)	(0.030)	(0.015)	(0.021)	(3.746E-04)	(5.240E-04)	(0.106)	(0.151)	(0.270)	(0.473)	(0.026)	(0.042)
Henan:														
Zhengzhou	6.867	-0.238	0.016	0.035	0.022	0.016	-2.130E-04	-2.397E-04	0.118	-0.003	-0.432	0.110	0.026	-0.005
	(0.108)	(0.231)	(0.008)	(0.016)	(0.006)	(0.011)	(1.459E-04)	(2.797E-04)	(0.044)	(0.083)	(0.139)	(0.334)	(0.011)	(0.025)
Kaifeng	6.512	0.306	0.028	0.019	0.031	-0.035	-4.433E-04	1.099E-03	0.219	-0.081	0.078	-0.508	-0.009	0.024
	(0.191)	(0.306)	(0.014)	(0.021)	(0.010)	(0.017)	(2.480E-04)	(4.132E-04)	(0.087)	(0.116)	(0.208)	(0.344)	(0.019)	(0.030)
Pingdingshan	6.800	-0.009	0.012	0.024	0.026	0.031	-3.020E-04	-7.686E-04	0.164	-0.085	-0.420	-0.808	0.023	0.059
	(0.202)	(0.284)	(0.015)	(0.021)	(0.012)	(0.015)	(3.001E-04)	(3.750E-04)	(0.087)	(0.116)	(0.235)	(0.339)	(0.021)	(0.029)
Xinxiang	6.498	0.335	0.017	0.014	0.044	-0.014	-6.083E-04	2.255E-04	0.141	0.022	-0.419	-0.104	0.033	-0.003
	(0.189)	(0.336)	(0.014)	(0.023)	(0.010)	(0.017)	(2.201E-04)	(4.025E-04)	(0.065)	(0.110)	(0.188)	(0.345)	(0.018)	(0.029)
Hubei:														
Wuhan	6.573	-0.055	0.027	0.029	0.043	0.000	-6.496E-04	8.910E-05	0.068	0.029	-0.287	0.053	0.020	-0.005
	(0.096)	(0.158)	(0.007)	(0.010)	(0.006)	(0.009)	(1.375E-04)	(2.220E-04)	(0.036)	(0.056)	(0.114)	(0.190)	(0.009)	(0.015)
Huangshi	6.663	0.377	0.011	0.035	0.055	-0.034	-8.971E-04	7.490E-04	0.038	0.146	-0.590	0.429	0.049	-0.036
	(0.150)	(0.279)	(0.011)	(0.018)	(0.007)	(0.015)	(1.746E-04)	(3.817E-04)	(0.055)	(0.084)	(0.156)	(0.307)	(0.014)	(0.027)
Guangdong:														
Guangzhou	7.328	0.196	0.010	0.032	0.024	0.025	-2.751E-04	-8.135E-04	0.109	0.069	-0.331	-0.141	0.019	0.013
	(0.144)	(0.255)	(0.011)	(0.018)	(0.007)	(0.012)	(1.725E-04)	(2.866E-04)	(0.058)	(0.088)	(0.158)	(0.273)	(0.014)	(0.023)
Foshan	7.834	0.315	-0.002	0.016	0.018	0.013	-2.386E-04	-3.678E-04	0.141	0.104	-0.365	-0.176	0.021	0.010
	(0.157)	(0.373)	(0.012)	(0.026)	(0.009)	(0.020)	(2.259E-04)	(4.850E-04)	(0.071)	(0.152)	(0.167)	(0.375)	(0.016)	(0.034)
Zhanjiang	6.708	1.155	0.019	-0.009	0.067	-0.030	-1.127E-03	2.795E-04	0.048	0.168	-0.253	-0.625	0.009	0.056
	(0.249)	(0.473)	(0.019)	(0.034)	(0.013)	(0.025)	(3.198E-04)	(6.392E-04)	(0.105)	(0.171)	(0.278)	(0.542)	(0.025)	(0.047)
Shenzhen	7.187	-1.018	0.023	0.028	0.084	0.021	-1.787E-03	6.920E-05	0.071	0.051	-0.219	-0.378	0.012	0.036
	(0.192)	(0.407)	(0.012)	(0.027)	(0.013)	(0.024)	(3.246E-04)	(5.706E-04)	(0.072)	(0.137)	(0.198)	(0.403)	(0.017)	(0.034)
Huizhou	6.941	0.454	0.028	-0.033	0.041	0.024	-6.931E-04	-5.614E-04	0.105	0.185	-0.393	-0.154	0.020	0.005
	(0.175)	(0.380)	(0.014)	(0.026)	(0.011)	(0.022)	(2.733E-04)	(5.165E-04)	(0.079)	(0.160)	(0.231)	(0.399)	(0.020)	(0.037)
Zhaoqing	6.738	0.950	0.024	-0.032	0.065	0.005	-1.106E-03	-2.392E-04	-0.018	0.151	-0.515	-0.559	0.033	0.050
	(0.206)	(0.456)	(0.016)	(0.032)	(0.012)	(0.024)	(2.967E-04)	(5.934E-04)	(0.096)	(0.162)	(0.230)	(0.524)	(0.022)	(0.044)

Table 1.A3: Estimates of Baseline Human Capital Model by Province and City (Continued)

	β_0	δ_0	β_1	δ_1	β_2	δ_2	β_3	δ_3	β_4	δ_4	β_5	δ_5	β_6	δ_6
Yunnan:														
Kunming	6.653	0.033	0.031	0.016	0.047	-0.005	-7.553E-04	2.133E-04	0.055	-0.060	0.005	-0.193	-0.004	0.015
	(0.084)	(0.184)	(0.006)	(0.012)	(0.005)	(0.010)	(1.306E-04)	(2.679E-04)	(0.033)	(0.065)	(0.092)	(0.209)	(0.008)	(0.017)
Gejiu	6.490	0.617	0.034	-0.010	0.059	-0.040	-9.269E-04	8.276E-04	-0.014	0.039	-0.422	0.051	0.027	0.004
	(0.116)	(0.248)	(0.009)	(0.016)	(0.007)	(0.015)	(1.532E-04)	(3.368E-04)	(0.047)	(0.077)	(0.127)	(0.238)	(0.012)	(0.021)
Dali	6.877	0.025	0.026	0.000	0.021	0.050	-1.345E-04	-1.285E-03	0.032	0.035	-0.245	-0.269	0.011	0.024
	(0.142)	(0.237)	(0.010)	(0.016)	(0.009)	(0.014)	(2.141E-04)	(3.309E-04)	(0.053)	(0.082)	(0.152)	(0.272)	(0.014)	(0.024)
Dongchuan	6.631	0.312	0.030	0.006	0.030	0.002	-2.473E-04	-1.455E-04	0.130	-0.067	-0.180	-0.507	0.006	0.044
	(0.191)	(0.305)	(0.013)	(0.020)	(0.011)	(0.016)	(2.405E-04)	(3.496E-04)	(0.071)	(0.094)	(0.208)	(0.296)	(0.019)	(0.026)
Baoshan	6.577	0.398	0.037	-0.007	0.032	0.000	-3.334E-04	-1.239E-04	-0.008	0.099	0.174	-0.256	-0.026	0.035
	(0.149)	(0.244)	(0.010)	(0.016)	(0.008)	(0.013)	(1.940E-04)	(3.113E-04)	(0.057)	(0.077)	(0.163)	(0.281)	(0.015)	(0.024)
Gansu:														
Lanzhou	6.467	-0.022	0.029	0.017	0.057	0.000	-8.713E-04	-7.210E-05	0.120	0.001	-0.654	-0.026	0.046	0.000
	(0.075)	(0.179)	(0.006)	(0.012)	(0.004)	(0.009)	(1.030E-04)	(2.264E-04)	(0.032)	(0.072)	(0.093)	(0.217)	(0.008)	(0.018)

Figure 1. Gender Gap in Log-Earnings by Education: 1988 and 1995



Note—The lines are predicted from models with a linear specification of schooling, and the markers are from the same models with a categorical specification of schooling. Experience is held at the sample average for each year.