Regional Variation in Earnings Inequality in Reform-Era Urban China[1]

Yu Xie and Emily Hannum
University of Michigan

This article studies the regional variation in earnings inequality in contemporary urban China, focusing on the relationship between the pace of economic reforms and earnings determination. Through a multilevel analysis, it shows that economic growth depresses the returns to education and work experience and does not affect the net differences between party members and nonmembers and between men and women. Overall earnings inequality remains low and only slightly correlated with economic growth because, in faster-growing cities, the tendency toward higher levels of inequality is somewhat offset by the lower returns to human capital. A plausible interpretation is that these results are largely due to the lack of a true labor market in urban China.

Post-1978 economic reforms have brought about rapid economic growth and high levels of personal income unprecedented in modern Chinese history. The per capita gross national product grew from 375 yuan in 1978 to 1,026 inflation-adjusted yuan in 1992, averaging an annual growth rate of 7.45%. During the same period, the per capita annual income for urban residents increased from 316 yuan to 721 inflation-

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adjusted yuan, an average annual rate of 6.07%.[2] Against the backdrop of an unstable economy and the Maoist egalitarian ethos in prereform China (Riskin 1987),[3] this drastic economic boost has spurred new interest among social scientists in understanding the consequences of economic reforms for income and wealth distribution in China (Trescott 1985; Adelman and Sunding 1987; Walder 1987, 1990, 1992*a*, 1992*b*; Riskin 1987; Hsiung and Putterman 1989; Nee 1989, 1991, 1994, 1996; Gelb 1990; Zhao 1990; Li 1991; Zhu 1991; Peng 1992; Khan et al. 1992; Griffin and Zhao 1993; Selden 1993; Xin 1994). These researchers have been intrigued by two fundamental questions: (1) Do economic reforms increase or reduce inequality? and (2) Do economic reforms disproportionately benefit certain social groups at the expense of others? Indeed, these two questions have dominated all theoretical discussions among sociologists of postsocialist or reforming-socialist economies. Such discussions were recently summarized by Róna-Tas (1994, table 1).[4]

With the exception of Khan et al. (1992), Knight and Song (1993), and Nee (1994, 1996), studies of the consequences of economic reforms in China typically treat China as a homogeneous entity and disregard enormous regional variations, even when data used are regional and unrepresentative of China as a whole. This practice, while sensible when Chinese data were scarce, should no longer be continued, as China is a vast country with spatially heterogeneous economies (Linge and Forbes 1990;

[2] Computed from statistics published by the State Statistical Bureau (1993, tables 2–12 and 8–6). All figures are in 1978 constant yuan. Although per capita annual income for rural residents grew much faster at an average rate of 8.86% per year during the same period, rural residents' average income was only 438 yuan per capita in 1992, about 40% lower than that for urban residents.

[3] As Whyte (1986) points out, inequalities in prereform China may have been substantial despite Maoist egalitarian rhetoric. In contrast with the preeminence of monetary inequality in capitalist societies, most significant forms of inequality in socialist economies lie in the distribution of goods and services, such as government housing, provided through government redistributive hierarchies (Szelényi 1978, 1983). Furthermore, although low by international standards, income inequalities were substantial in prereform China, largely due to the enormous gap between rural and urban residents (Whyte 1986; Adelman and Sunding 1987).

[4] Although Róna-Tas's summary appears succinct and elegant, his characterization of Nee's (1989, 1991) theory as implying that "inequalities should not increase" (Róna-Tas, p. 43) due to "structural compensation" between market and bureaucratic coordinations (p. 47) appears misleading. Nee's market transition theory essentially argues that economic reforms in state socialism gradually replace bureaucratic redistributive coordination with market coordination and thus favor producers with productive skills over government bureaucrats. Although Nee (1991, p. 269) is clearly aware of new sources of inequalities generated by markets, he is vague on the trend in overall inequality during economic reforms. As is shown later in this article, Nee's theory actually implies that overall inequality increases as returns to education increase.

Li 1993). Ignoring China's spatial heterogeneity during economic reforms is particularly striking in light of the fact that "dozens of books and hundreds of journal articles" have focused on the reform in Hungary alone (Kornai 1989, p. 32), a country smaller in population size than 25 of China's 30 province-level administrative units.[5]

Consideration of regional heterogeneity is significant, not only because economic activities in different parts of China are dictated by large regional variations in natural and human resources, but also, and more important, because the Chinese industrial reform has had a regional dimension. Aguignier (1988), Falkenheim (1988), Shirk (1989), Linge and Forbes (1990), and Li (1991, 1993) all document that the reform has disproportionately benefited coastal provinces at the expense of inland provinces and that serious tensions across regions have resulted. What has often been overlooked, however, is that the increasing regional disparities were part of a *deliberate* scheme from the outset. There are several reasons for this. First, the central government thought it was "less dangerous . . . to carry out the initial experiments in a distant province [such as Guangdong] rather than in Tianjin, Shanghai, or elsewhere in the country's industrial heartland" (Linge and Forbes 1990, p. 15). Since the experiments were contained in restricted regions, their failure would not have been catastrophic to the national economy. Second, following a long-standing practice in prereform China, Chinese reformers wished to set up a few exemplary models to showcase the industrial reform. Channeling government resources and foreign investment to limited regions would greatly enhance the likelihood of success. At the core of the Chinese Communist Party's plan for economic reforms has been the slogan, "Let certain people become rich first in order to achieve common prosperity" (Zhao 1994, p. 115). It is hoped that the affluence of the coastal regions will trickle down to the remote hinterlands, as the whole country starts to emulate the success of the coastal regions. Finally, Chinese leaders learned the value of decentralization from the agricultural reform and encouraged it during the early stages of the industrial reform, allowing local governments to use their comparative advantages and to "sidestep their 'shortcomings'" (Falkenheim 1988, p. 287). This permission to use comparative advantage was a primary reason for the rise of Shenzhen, near the border between Guangdong and Hong Kong, from a small town to a sizable metropolis with a vibrant economy and a major stock market.

With data from a 1988 national income survey of 9,009 urban households, this article studies the regional variation in earnings inequality in

[5] The exceptions are Tianjin, Hainan (newly separated from Guangdong), Tibet, Qinghai, and Ningxia.

China. Capitalizing on the fact that the pace of economic reforms has been regionally uneven, we examine the relationship between the success of economic reforms, measured by city-level economic growth between 1985 and 1988, and individual-level earnings determination. Our analysis consists of four steps. First, we develop and estimate a modified human capital model that takes into account political advantages important in the Chinese context. Second, we decompose employment earnings into regular salary or wage and cash bonuses and subsidies. Third, we consider multilevel models of regional heterogeneity with parameters of the baseline human capital model determined by an indicator of economic growth across cities. Last, we draw inferences about the regional variation in overall inequality and its relation to economic growth.

A BASELINE HUMAN CAPITAL MODEL

We modify Mincer's (1974) human capital model for contemporary China into the form of

$$\begin{aligned} T = \log Y = \beta_0 + \beta_1 X_1 + \beta_2 X_2 + \beta_3 X_2^2 \\ + \beta_4 X_4 + \beta_5 X_5 + \beta_6 X_1 X_5 + \epsilon, \end{aligned} \tag{1}$$

where Y is earnings, X_1 years of schooling, X_2 years of work experience, X_4 a dummy variable denoting membership in the Communist Party of China (1 = party member), and X_5 a dummy variable denoting gender (1 = female). All β's are unknown parameters, and ϵ is the residual unexplained by the model. Equation (1) deviates from Mincer's model in two ways. First, we include party membership in the model and interpret it as an aspect of human capital associated with political advantages.[6] Past research (e.g., Walder 1990; Knight and Song 1993) clearly documents the importance of party membership and lends support to our interpretation. Second, we apply the model to both male and female workers and allow for differences between the sexes in the intercept as well as in the return to years of schooling. Inclusion of women into the earnings equation improves upon similar models based on data only from male workers for the United States and other countries. The common practice of excluding women often stems from the difficulty of dealing

[6] We intend party membership to capture Nee's (1989, 1991, 1994) notion of "political capital" or "positional power" that should decline in importance relative to "market capital" or "performance" during economic reforms. For this purpose, our measure is less ideal than Nee's various types of "cadres." As pointed out by James Heckman at the 1994 IRP workshop, party membership can be interpreted as part of human capital, as individuals may purposefully invest in being a party member in order to reap monetary benefits.

with women not currently participating in the labor force. For contemporary China, however, this difficulty does not arise as women's labor force participation is nearly universal. We will discuss the justification for including the interaction between gender and education later.

For estimation, we use data from the 1988 Chinese Household Income Project (CHIP). CHIP encompassed two surveys, one for urban residents and another for rural residents. For this article, we use data only from the urban portion of the study. The urban survey follows a multistage sampling methodology, with the first step being to select 10 province-level administrative units (out of a total of 30) and then 55 cities (out of a total of 434) to represent varying urban conditions in China (Eichen and Zhang 1993). The urban survey instrument was administered to a total of 9,009 households in March and April 1989.

The survey gathered information pertaining to all household members, including demographic and educational characteristics and labor force activities. For this analysis, we make an unrealistic but convenient assumption that within-household clustering is negligibly small. That is, we treat all members of sampled households who are between ages 20 and 59 and active in the labor force as independent observations. Combined with other criteria for excluding respondents with missing or incomplete data, this procedure yields a sample of 15,862 cases.[7]

As a study of income, the CHIP survey instruments were carefully tailored to capture all forms of income in 1988, including the provision of cash bonuses and subsidies. Major findings from the study are published in a book edited by Griffin and Zhao (1993). In this article, we are concerned with the cash compensation that individual workers earn at the workplace. Specifically, our measure of earnings (Y) contains the following three components: (*a*) (regular monthly salary/wage) $\times$ 12, denoted by Y_1; (*b*) (monthly average of cash bonuses and subsidies) $\times$ 12, denoted by Y_2; and (*c*) 1988 annual earnings from private businesses, denoted by Y_3. The last component plays a negligible role (about 1% of total earnings). Between the first two forms of earnings, regular salary/wage (Y_1) is slightly more important.[8]

Following Mincer (1974, p. 48), we extrapolate years of schooling from levels of attained education (less then three years of schooling = 1; three

[7] Since our study mainly focuses on regional variation in earnings determination as a function of economic growth, we excluded respondents from counties for which we have no reliable macrolevel data. This criterion resulted in the exclusion of 926 respondents from our study. Through additional sensitivity analysis (unreported here but available upon request), we confirmed that our results for the first part of the study are unaffected by this exclusion criterion.

[8] In the aggregate, bonuses and subsidies comprise about 45% of total employment earnings (i.e., after excluding earnings from private businesses). See app. table A1.

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). Given the relatively unfamiliar and poorly understood setting of contemporary China, the conservative treatment of education as a truly categorical variable (in levels of attained education) has been wisely adopted by many researchers (e.g., Peng 1992; Khan et al. 1992; Griffin and Zhao 1993; Nee 1994, 1996). However, an interval measure in years of schooling is preferred for our study due to three important considerations. First, the theoretical framework of human capital requires that education be considered in years of schooling as a primary source and thus cost of investment (Mincer 1974). Second, with log of earnings as the dependent variable as in equation (1), the coefficient for years of schooling can be interpreted readily as the rate of return and compared cross-nationally (Psacharopoulos 1981). Finally, a one-degree-of-freedom specification for education effects allows us to conveniently analyze geographic variation in returns to schooling in relation to the local economic context. Even if the stringent linearity specification is imperfect, it may be a good approximation. To check the suitability of this simplifying specification, we will conduct a sensitivity analysis.

We also follow Mincer (1974, p. 48) in calculating years of work experience as the difference between the current age and the age at first year of experience, 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). Again, this parameterization is dictated by the choice to build our baseline model on human capital theory, which requires that "it is experience rather than age that enters as an independent variable (in order to distinguish between the biological and human capital effects of time)" (Psacharopoulos 1977, p. 40).

The first column of table 1 presents ordinary least squares (OLS) estimates of equation (1) with the constraint $\beta_6 = 0$. Thus, model 1 forces the effect of schooling to be parallel between men and women. With this rather restricted model, we begin to draw some general inferences about the earnings determination in urban China. The negative estimates for β_3 confirm the expectation from human capital theory that the experience effect should be concave, first increasing for most of the working life and then declining toward the end of the working life (Mincer 1974, p. 84). According to model 2, the optimal level of experience is 33.2 years, a figure close to the 33.8 years implied by Mincer's (1974, p. 92) estimates for the United States.[9]

[9] The optimal number of years of experience is obtained by solving $\partial \log Y/\partial X_2 = 0$.

TABLE 1

THREE REGRESSION MODELS FOR TOTAL EARNINGS ASSUMING REGIONAL HOMOGENEITY

INDEPENDENT VARIABLE	MODEL 1		MODEL 2		MODEL 3	
	Parameter	SE	Parameter	SE	Parameter	SE
Intercept (β_0)	6.591	.017	6.685	.019	6.870	.017
Years of schooling (β_1)	.031	.001	.022	.001		
Level of education:*						
Junior high school					−.008	.015
Senior high school					.071	.016
Technical school					.082	.018
Community college					.137	.020
Four-year college and above					.226	.019
Experience (β_2)	.044	.001	.046	.001	.047	.001
Experience^2 (β_3)	$(-6.63)10^{-4}$	$(2.54)10^{-5}$	$(-6.93)10^{-4}$	$(2.54)10^{-5}$	$(-7.25)10^{-4}$	$(2.61)10^{-5}$
Party member (1 = yes) (β_4)	.071	.008	.073	.008	.074	.008
Gender (1 = female) (β_5)	−.114	.006	−.344	.021	−.302	.017
Gender × years of schooling (β_6)			.022	.002		
Interaction of gender and level of education:						
Junior high school					.173	.019
Senior high school					.217	.021
Technical school					.265	.024
Community college					.281	.029
Four-year college and above					.272	.031
Sum of squares error	2,179.2		2,161.8		2,160.0	
df	15,856		15,855		15,847	
R^2 (%)	26.14		26.73		26.79	

NOTE.—N = 15,862. The dependent variable (T) is the natural logarithm of total annual earnings (yuan). β's refer to ordinary least squares estimates of eq. (1).
* Excluded = primary or less.

According to model 1, the estimated return in earnings to years of schooling is 3.1%. By international standards, this figure is extremely small. In his review with an emphasis on international comparison, Psacharopoulos (1981, p. 330) places the rate of return to education between 5.9% (for Canada) and 22.8% (for Malaysia). The very low estimated rate of return for China is particularly puzzling in light of China's status as a less developed country and its rapid economic growth since 1978. Both of these traits are commonly associated with high returns to education. Psacharopoulos (1981) reports that the rate of return tends to be higher in less developed countries (with an average of 14.4%) than in more economically advanced countries (with an average of 7.7%). Economic reasoning also predicts that the return to schooling is positively related to the rate of economic growth because "individuals who are more efficient resource allocators will be better able to take advantage of the changed opportunity sets" (Chiswick 1971, p. 28).

We have no reason to doubt the reliability of the CHIP data in light of the low estimated return to education, for many other studies have also found similarly low rates of return to education in China using independent data sources. From a sample drawn in Nanjing, for example, Byron and Manaloto (1990, p. 790) report an "astonishingly low rate of return of income to education, about 4% for each additional year of schooling."[10] From a Tianjin sample, Walder (1990, pp. 149–50) estimates a 1.0% rate of return in income and a 1.6% rate of return in salary. In addition, earlier research on selected Chinese subpopulations found education to have either no effect (e.g., Whyte and Parish 1984; Zhu 1991) or negative effects (e.g., Gelb 1990; Peng 1992; Nee 1994) on income. Thus, we are reassured by the consistency of the low rate of return to education estimated in model 1 with the relevant literature.

Coefficient β_4 reveals the advantage of being a member of the Communist Party of China. According to model 1, party members earn about 7.4% more than nonmembers net of education, experience, and gender. This estimate is close to the 9% party premium reported by Walder (1990, p. 150) for the 1976 and 1986 salaries of a Tianjin sample. As cautioned by Knight and Song (1993, pp. 253–59), however, the effect of party membership may not be causal. The party may selectively recruit workers possessing characteristics associated with high productivity, some of which may be unobserved and thus unconsidered here. In any event, our interest in using party membership is to detect the regional variation in party membership effects. Unless the potential selectivity varies substan-

[10] In fact, Byron and Manaloto's estimate was inflated after correcting for measurement errors. Before the correction, the estimate varies from 1.2% to 1.9% depending on estimation methods (p. 788).

tially across regions, using party membership as an indicator of positional power should serve our purpose well.[11]

The estimate for β_5 in model 1 indicates that women on average earn about 10.8% less than men of equal education, experience, and party status. This estimate is consistent with those reported by Byron and Manaloto (1990), Gelb (1990), Peng (1992), Walder (1990), and Knight and Song (1993), which range from 5 to 14%. Although the gender gap is relatively small by international standards,[12] it nonetheless exists and is substantial despite the Chinese government's rhetoric on gender equality dating from the founding of the People's Republic of China in 1949 (for a review, see Hannum and Xie 1994). The gender gap in earnings, however, is not the same across the spectrum of educational attainment. In model 2, we test the interaction between gender and schooling.

Model 2 allows β_6 to be free and thus contains all of the parameters specified by equation (1). The F-test statistic from nesting models 1 and 2 is 128 with 1/15,855 degrees of freedom, highly significant, indicating that the schooling effect is not parallel between men and women. The positive interaction effect for gender and schooling reveals that the return to schooling is higher for women than for men. Estimates of model 2 indicate that women's return to schooling is 4.5%, twice that of men's (2.2%). The interaction effect between gender and schooling is essentially a consequence of women's significantly lower earnings at low levels of education.[13] This can be easily seen in figure 1, where the two solid lines represent the predicted log of earnings as a function of years of schooling, separated for men and women using the estimates of model 2.

To guard against a possible misspecification of the education effect through the linearity constraint imposed on models 1 and 2, we conduct a sensitivity analysis in model 3, which allows for the effect of education to be distinct at six levels through the use of discrete coding. The F-test statistic for nesting models 2 and 3 is 1.65 with 8/15,847 degrees of freedom, not significant at 0.05 level. Thus, the linear specification for education effects seems to be an acceptable approximation. Figure 1 contrasts predicted earnings by schooling and gender according to models 2 and 3 after controlling for other independent variables. The figure shows that a linear specification of education leads to a slight underpre-

[11] In other words, one needs to entertain the possibility of three-way interactions among party membership, unobserved characteristics, and region in order to question our use of party membership. As argued elsewhere (Xie 1989), such high-order interactions are best assumed to be absent if there is neither good theory nor strong evidence in their support.

[12] In the United States, e.g., the gender gap in earnings net of education and age remained above 30% between 1960 and 1980 (Bianchi and Spain 1986, p. 177).

[13] We thank William Parish for pointing this out to us.

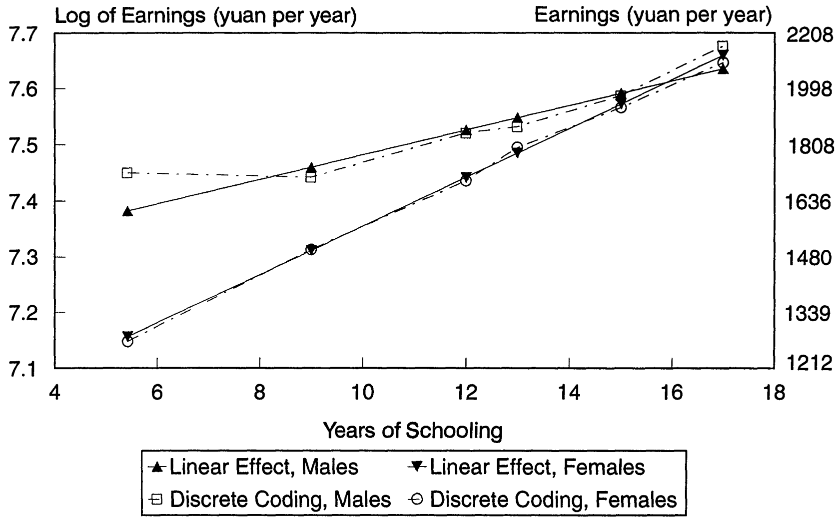

FIG. 1.—Education effects on earnings in China, comparison of two specifications (based on models 2 and 3 in table 1; other variables are held at sample means).

diction for the least-educated men (with primary or less education). Aside from this, the linearity specification proves a good approximation. Hence, we will use the linear specification in the remainder of this article.

DECOMPOSING TOTAL EARNINGS

The dependent variable for the regression models in table 1 is $T = \log Y$, where $Y = (Y_1 + Y_2 + Y_3)$. Recall that Y_1, Y_2, and Y_3 respectively represent (*a*) regular salary/wage, (*b*) cash bonuses and subsidies, and (*c*) earnings from private enterprises. Since only a negligibly small proportion of Chinese urban residents had any substantial earnings from private enterprises in 1988,

$$T = \log Y \approx \log(Y_1 + Y_2).$$

As noted by Knight and Song (1993), it is important to further decompose the total earnings (Y) into its two principal components (Y_1 and Y_2), as the two components are quite distinct from each other. Before economic reforms, earnings in urban China consisted primarily of Y_1. With the progress of the economic reforms, Y_2 has gained more significance. Thus, we reason that the impact of economic reforms on earnings is best manifested through Y_2.

An obvious approach to separating Y_1 and Y_2 is to run models with

$\log(Y_1)$ and $\log(Y_2)$ as dependent variables. In fact, this strategy was adopted by Knight and Song (1993). However, we see two major disadvantages with this approach. First, $\log(Y_1 + Y_2)$ cannot be decomposed into linear functions of $\log(Y_1)$ and $\log(Y_2)$, and as a result regressions using $\log(Y_1)$ and $\log(Y_2)$ as dependent variables do not have a simple relationship to regressions using $\log Y$. Second, $\log(Y_1)$ is undefined for zero Y_1, and $\log(Y_2)$ for zero Y_2. One may argue that a respondent who does not have a regular salary or wage is not really employed and should be excluded from an analysis of employment-related earnings. The same argument does not hold for respondents with zero Y_2 as it is logically possible (albeit empirically unlikely) that some workers receive no other forms of compensation aside from regular salary or wage.[14] Recoding zero values of Y_2 to a small positive number is also problematic, for regression results are sensitive to the arbitrary choice of the small number.

To solve both problems, we devised the following decomposition method:

$$\begin{aligned}\log(Y_1 + Y_2) &= \log(Y_1) + \log(1 + Y_2/Y_1)\\ &= \log(Y_1) - \log[Y_1/(Y_1 + Y_2)] \qquad (2)\\ &= S + B.\end{aligned}$$

We define $S = \log(Y_1)$ as our dependent variable in regressions modeling regular salary/wage, and $B = -\log[Y_1/(Y_1 + Y_2)]$ as our dependent variable in regressions modeling the share of bonuses and subsidies relative to regular salary or wage. For the decomposition of equation (2) to work, Y_1 must be a positive number. We delete respondents with no regular salary or wage and thus focus exclusively on *employed* respondents in our decompositional analyses. However, Y_2 is free to vary from zero to any positive number. When Y_2 is zero, B is also zero. Note that B does not measure the absolute amount of bonuses and subsidies but the relative share of bonuses and subsidies in comparison to the regular salary or wage. It is easy to show that B is closely related to bonus rate, defined as $Y_2/(Y_1 + Y_2)$:[15]

$$B = -\log[Y_1/(Y_1 + Y_2)] = -\log[1 - Y_2/(Y_1 + Y_2)]. \qquad (3)$$

For our data set, the correlation between B and $Y_2/(Y_1 + Y_2)$ is 0.960.

[14] There are 191 such respondents in the data used for the previous regression models.

[15] This expression is similar to the complementary log log transformation widely used in generalized linear models in the form of log $[-\log(1 - r)]$, where r is usually a rate. We do not take the logarithm of B for two reasons: (*a*) B is easily interpretable as a linear part of $\log Y$ and thus has the same scale as S ($\log Y_1$); and (*b*) we expect B to be zero for some respondents.

TABLE 2

REGRESSION MODELS FOR DECOMPOSING EMPLOYMENT EARNINGS INTO SALARY/WAGE AND BONUS SHARE

INDEPENDENT VARIABLE	SALARY/WAGE (S)		BONUS SHARE (B)	
	Parameter	SE	Parameter	SE
Intercept (β_0)	6.039	.014	.646	.014
Years of schooling (β_1)	.029	.001	−.007	.001
Experience (β_2)	.037	.001	.008	.001
Experience^2 (β_3)	$(-3.74)10^{-4}$	$(1.87)10^{-5}$	$(-3.02)10^{-4}$	$(1.94)10^{-5}$
Party member (1 = yes) (β_4)	.075	.006	−.003	.006
Gender (1 = female) (β_5)	−.225	.016	−.098	.016
Gender × years of schooling (β_6)	.013	.001	.007	.001
Sum of squares error	1,124.0		1,213.2	
df	15,581		15,581	
R^2 (%)	44.60		3.27	

NOTE.—N = 15,588. Respondents with no regular salary are excluded from the analysis. The dependent variables are defined as follows: $S = \log(Y_1)$, and $B = -\log[Y_1/(Y_1 + Y_2)]$, where Y_1 denotes regular salary/wage and Y_2 denotes cash bonuses/subsidies.

Equation (3) also shows that B is a monotonic transformation of Y_2 for any fixed Y_1. For the convenience of decomposing log Y, we use S and B as our dependent variables and interpret B as the "bonus share" among total employment-related earnings. Variation in B essentially indicates the variation in the significance of bonuses and subsidies relative to the regular salary or wage. If B were constant, bonuses and subsidies would comprise the same proportion among total employment earnings.

We now apply the baseline model in the form of equation (1) separately to the two components of log Y, excluding respondents with no regular salary or wage. The regression estimates are shown in table 2. Note that parameter estimates are in the same scale across the two equations for each of the independent variables. Summing them for each row would give the net effect on $\log(Y_1 + Y_2)$, a quantity close to T. It is thus feasible for us to compare the estimated coefficients for the S and B equations in table 2 to those for the T equation of model 2 in table 1. Several interesting results emerge. First, the baseline model does a much better job in determining regular salary or wage than determining bonus share. The R^2 is 44.60% for the salary/wage (S) equation but only 3.27% for the bonus share (B) equation. Second, the effect of party membership (β_4) resides entirely in the determination of salary/wage. This is easy to see in the insignificant coefficient of party membership for B and in the comparable magnitude of the coefficients for S and T (0.075 and 0.073, respectively). Third, by the same logic, experience affects total earnings primarily through salary/wage. About four-fifths of the linear effect of experience (β_2) works through salary/wage, and the rest works through bonus share. Fourth, more than two-thirds of the gender gap for those with no schooling (β_5) is due to men's higher salaries/wages (S), and the rest is due to men's higher share of bonuses and subsidies (B). Fifth, the pattern of education effects differs between men and women. For men, the return to schooling is *suppressed* by the bonus share, as the return is negative (-0.007) for B and thus higher for S (0.029) than for T (0.022). For women, schooling has no significant effect on the share of bonuses and subsidies. Combined, these results indicate that a large portion (about one-third) of the gender-schooling interaction effect reported earlier for T (0.022) is due to bonuses and subsidies rather than to regular salaries/wages, as men's return for B is negative.

MODELING REGIONAL HETEROGENEITY

The Model

Although the application of the baseline human capital model of equation (1) is a useful first step, treating China as a homogeneous entity is both methodologically untenable and theoretically wasteful. It is methodologically untenable because regional variation in earnings determination is substantial, as is shown in appendix table A2. It is theoretically wasteful because we can capitalize on the regional variation in earnings determination to test theories relating earnings inequality to the progress of economic reforms. Thus, we relax the assumption that the human capital model in the form of equation (1) is regionally homogeneous.

Taking regional heterogeneity to an extreme, one may treat different cities as totally different regimes and allow full interactions between cities and the coefficients of equation (1). Results from this exercise are presented in appendix table A2. While one may gain some insight by examining city-specific estimated coefficients, this strategy is unduly conservative and does not allow us to test theoretically interesting hypotheses. For example, we would like to know whether the regional variation in earnings determination can be explained by the regional variation in economic growth. To this end, we have developed the following multilevel model with two components, one at the individual level and another at the city level. Our strategy is similar to that of DiPrete and Grusky (1990), except that our macrolevel variation is regional rather than temporal.

Using total earnings (T) again as an illustration, for the ith person ($i = 1, \ldots n_k$) in the kth city, the model at the individual level is

$$\log(y_{ik}) = \beta_{0k} + \beta_{1k}x_{1ik} + \beta_{2k}x_{2ik} + \beta_3 x_{2ik}^2 + \beta_{4k}x_{4ik} + \beta_{5k}x_{5ik} + \beta_6 x_{1ik}x_{5ik} + \epsilon_{ik}. \tag{4}$$

A distinct feature of equation (4), as compared to equation (1), is that the coefficients of x_1, x_2, x_4, and x_5 (i.e., β_{0k}, β_{1k}, β_{2k}, β_{4k}, and β_{5k}) vary across cities. Note that a constraint is placed on β_3 and β_6 so that they do not change with k.[16] At the city level, we assume

$$\beta_{0k} = \alpha_0 + \lambda_0 z_k + \mu_{0k}, \tag{5a}$$

$$\beta_{1k} = \alpha_1 + \lambda_1 z_k + \mu_{1k}, \tag{5b}$$

$$\beta_{2k} = \alpha_2 + \lambda_2 z_k + \mu_{2k}, \tag{5c}$$

$$\beta_3 = \alpha_3, \tag{5d}$$

$$\beta_{4k} = \alpha_4 + \lambda_4 z_k + \mu_{4k}, \tag{5e}$$

$$\beta_{5k} = \alpha_5 + \lambda_5 z_k + \mu_{5k}, \tag{5f}$$

$$\beta_6 = \alpha_6. \tag{5g}$$

The city-level variable z measures economic growth. For this study, we compute z by

$$z = \log(\mathrm{GPVI}_{1988}/\mathrm{GPVI}_{1985}), \tag{6}$$

where GPVI stands for the gross product value of industry. For each city, we extracted GPVI data from published tables in *China Urban Statistics* (*CUS*) 1985 and 1988 (State Statistical Bureau 1985, 1990).[17] Despite the popularity of energy consumption as a measure of modernization in cross-national studies (Crenshaw and Ameen 1994, p. 10), we chose to use the gross product value of industry to measure economic growth directly. Energy consumption would be inappropriate for our study, as it is highly sensitive to the composition of local industries (i.e., light vs. heavy industries). Note that

$$z = \log(\mathrm{GPVI}_{1988}) - \log(\mathrm{GPVI}_{1985}).$$

[16] This restraint is tantamount to forcing the effects of education and experience either to increase or to decrease across cities and thus greatly facilitates interpretation and hypothesis testing.

[17] Besides the gross product value of industry, we experimented with other economic indicators (such as the net output value and the number of enterprises) available in the same data source. We decided to use the gross product value of industry because it is least likely to be misreported while being indicative of real economic growth. As Victor Nee pointed out to us in a telephone conversation on May 19, 1995, economic growth does not measure institutional changes in property rights and labor markets that may have taken place since 1989.

That is, z measures the *change* in economic output, or economic *growth,* between 1985 and 1988, net of regionally fixed characteristics associated with economic output such as the size, infrastructure, and natural and social conditions of local economies (Firebaugh and Beck 1994, p. 636). In addition, z can be interpreted as a monotonic transformation of the more familiar growth rate (denoted as r):

$$r = (\mathrm{GPVI}_{1988} - \mathrm{GPVI}_{1985})/\mathrm{GPVI}_{1985} = \exp(z) - 1.$$

In appendix table A1, we present z and an annualized growth rate, along with other city-based descriptive statistics computed from the CHIP data set.[18] We note that z varies widely among the 55 cities, ranging from .19 (corresponding to a 6.54% annual growth rate) for Yangquan of Shanxi Province to 1.16 (corresponding to a 47.21% annual growth rate) for Shenzhen of Guangdong Province. As expected, the growth rate tends to be higher in the coastal provinces of Guangdong and Jiangsu than in other provinces.

Our interest centers on how the determinants of earnings vary as a function of economic growth in China. This is made clearer if we substitute (5) into (4):

$$\begin{aligned}\log(\hat{y}_{ik}) = {} & \alpha_0 + \alpha_1 x_{1ik} + \alpha_2 x_{2ik} + \alpha_3 x_{2ik}^2 + \alpha_4 x_{4ik} + \alpha_5 x_{5ik} + \alpha_6 x_{1ik} x_{5ik} \\ & + \lambda_0 z_k + \lambda_1 x_{1ik} z_k + \lambda_2 x_{2ik} z_k + \lambda_4 x_{4ik} z_k + \lambda_5 x_{5ik} z_k \qquad (7) \\ & + (\mu_{0k} + \mu_{1k} x_{1ik} + \mu_{2k} x_{2ik} + \mu_{4k} x_{4ik} + \mu_{5k} x_{5ik} + \epsilon_{ik}).\end{aligned}$$

The second line of equation (7) represents the interactions of the individual-level explanatory variables (including the intercept but excluding squared work experience and gender-education interaction) and the city-level measure of economic growth. The third line of equation (7) indicates the composite residual consisting of an individual-level residual and city-level residuals weighted by individual-level independent variables.

A general two-level model, equation (7) provides the framework for our analysis of the regional variation in earnings determination (see Mason, Wong, and Entwisle 1983). Some special cases of this model are worth mentioning:

[18] These figures have been adjusted for inflation. While geographic boundaries could have changed between 1985 and 1988, we do not adjust for population changes in these years for two reasons. First, we want our measure to be able to capture real economic growth associated with urbanization over time. Second, it is not clear that the State Statistical Bureau consolidated their data on population with those on industrial output in terms of geographic coverage. To the extent that both accounting procedures were subject to error, adjusting for population size would introduce additional errors. Since our GPVI measure includes gross product value of industry in a city's county or counties, reclassification of city boundaries should not be a serious problem.

A. If all λ's are zero, the model becomes the "random coefficients" model. In this case, economic growth does not have a systematic impact on the influence of the earnings determinants, whose effects vary randomly across cities.
B. If all λ's are zero and μ_{1k}, μ_{2k}, μ_{4k}, and μ_{5k} are also zero, the model becomes the "variance components" model. In this case, only the overall level of earnings (indicated by the intercept) varies randomly across cities.
C. If all λ's and all μ's are zero, β's are fixed across k. This model reduces to an individual-level model and can be estimated via OLS. In this case, intercity variation is ignored, and regional homogeneity is assumed. Such models were presented in table 1.

HYPOTHESES

One interpretation of our measure of economic growth, z, is to treat it as an indicator of the success of urban economic reforms between 1985 and 1988. Note that z is not an "intention" measure but an "outcome" measure. If the urban economic reforms have achieved what they were intended to do, namely to spur economic growth, the distinction is of minor significance. To the extent that efforts at urban economic reform have not always been successful (e.g., Shirk 1989; Walder 1992*a*, 1992*b*), the distinction is important, for, with the use of z, we are actually assessing the consequences for earnings determination of the *success* of economic reforms rather than of economic reforms themselves. With this caveat, we formulate the following two hypotheses.

HYPOTHESIS 1.—*The faster the economic growth, the greater the rate of return to education.*

We derive this hypothesis both from Nee's (1989) market transition theory and from Chiswick's (1971) economic analysis. According to Nee's market transition theory, during economic reforms, markets gradually replace state bureaucracies in reallocating surplus. Since markets "[favor] direct producers relative to redistributors" and accordingly reward productivity instead of political loyalty, the transition to a market economy is likely to give rise to "higher returns of education, which is among the best indicators of human productivity" (p. 666). If we accept our measure of economic growth (z) as a valid indicator of the success of economic reforms and thus market transition, we expect economic growth to be positively related with the returns to education.

In fact, this prediction preceded Nee's work on China. In an economic analysis of the relationship between earnings inequality and economic development, Chiswick (1971) essentially came to the same conclusion. Although he does not clearly identify a relationship between the rate of return and the *level* of economic development, Chiswick predicts that "the rate of return is likely to be positively related to the secular rate of *growth* of output" (p. 27; emphasis added). Chiswick's explanation is that economic growth associated with increased productivity and improved technology tends to provide better opportunities for the educated labor force, who disproportionately contribute to and derive benefits from a fast-growing economy. Thus, this economic reasoning leads to the expectation that our measure of economic growth (z) is positively related to the rate of return to education. That is, λ_1 of equation (7) should be positive.

HYPOTHESIS 2.—*The faster the economic growth, the smaller the returns to party membership.*

Based on the argument that market forces gradually replace political hierarchies in determining social stratification, Nee (1989) unambiguously projects a "decline in the value of political capital" (p. 671) during market transition. This is made possible, according to Nee's market transition theory, by the public's increasing reliance on market coordination and decreasing reliance on state functionaries for economic goods and services. For support, Nee uses Walder's findings from a Tianjin survey that the positive effect of party membership declined between 1976 and 1985 (see Walder 1990). Further, Nee suggests that the extent to which a socialist economy is replaced by a market economy can vary substantially by region and sector and that such variations have direct implications for social stratification (1989, p. 667). In our analysis, we operationalize Nee's suggestion through a multilevel model with a macrolevel z variable measuring the success of market transition. If Nee's prediction is correct, we expect to see a negative coefficient for λ_4.

Besides these two hypotheses, we are interested in the interaction effects between economic growth and work experience (λ_2) and between economic growth and gender (λ_5), although our theoretical expectations for these interactions are not clear-cut. Work experience may be more important in the old redistributive economy because experience is built into bureaucratic formulas for setting salaries under the old regime but is subject to market regulations under the new regime. If experience boosts productivity by an amount that exceeds the experience schedule for pay formally set by bureaucratic formulas, experience should be positively related to economic growth. However, we have no good reason to endorse this proposition, especially in light of the fact that experience may become quickly obsolete in a fast-growing economy. Thus, we postpone theoretical discussions about the relationship between economic growth and the returns to work experience until empirical results are presented.

In a similar fashion, we are reluctant to form an unequivocal hypothesis concerning the relationship between the gender gap in earnings and economic growth. Fears that economic reforms will exacerbate gender inequalities have been voiced (e.g., Trescott 1985). Clearly, the main source of the fears lies in the erosion of the government's ability to implement socialist affirmative-action-type programs aimed at reducing gender inequalities. In other words, if gender inequalities are held to a low level by government interventions, they may climb during the transition to markets. However, Walder (1990) challenges this speculation and reports evidence instead for the narrowing of the gender gap in income.

We follow our earlier strategy of decomposing the logarithm of total earnings (T) into a regular salary/wage component (S) and a bonus/subsidy share component (B) within the multilevel model framework. Not only are we interested in testing hypotheses concerning how economic growth influences the effects of schooling and party membership on total earnings, we would also like to know if the influence of economic growth works through regular salaries and wages or through bonuses and subsidies. Even when economic growth is unrelated to the effects of individual-level determinants on total earnings, it may influence the effects of the individual-level determinants on the regular salary/wage and the bonus/subsidy share in different directions.

Results

To estimate the multilevel model of equation (7), we make the following assumptions. First, our explanatory variables are exogenous, that is, all x's and z are uncorrelated with ϵ and μ. Second, ϵ_{ik} is independent across both the i and k subscripts and follows an identical distribution with an expected value of zero and a variance of σ_ϵ^2. Third, residuals are uncorrelated across levels, that is, $\text{cov}(\epsilon, \mu_p) = 0$ (for $p = 0, 1, 2, 4, 5$). Fourth, city-level residuals have a joint distribution with a mean of 0 and a diagonal variance-covariance matrix Ω_μ.[19] Given this set of assumptions, the model can be estimated via an iterative generalized least squares (IGLS) method, which updates the estimation of the variance-covariance matrix of the composite residual (i.e., the third line marked by [b] in eq. [7]). As Goldstein (1986) shows, this IGLS method is identical to maximum-likelihood estimation under the assumption that the random components of the model (ϵ_{ik} and μ_p's) follow a multivariate normal

[19] The assumption of zero covariances among city-level residuals is not required for estimation. We did estimate many multilevel models with nonzero covariances, but we do not allow nonzero covariances here because they are empirically insignificant. Our presentation of the results is greatly simplified by this restriction.

TABLE 3

GOODNESS-OF-FIT STATISTICS FOR SELECTED MULTILEVEL MODELS

Model Specification	No. of Parameters	L^2	χ^2	*df*
Total earnings (T) (N = 15,862):				
Variance components model	9	10,132.2		
Random coefficients model	13	10,016.5	115.7*	4
Full model	18	9,986.2	30.3*	5
Salary/wage (S) (N = 15,588):				
Variance components model	9	1,126.6		
Random coefficients model	13	975.4	151.2*	4
Full model	18	955.1	20.3*	5
Bonus share (B) (N = 15,588):				
Variance components model	9	1,003.2		
Random coefficients model	13	966.9	36.3*	4
Full model	18	926.7	40.2*	5

NOTE.—Covariances among macrolevel error terms are constrained to be zero. No. of parameters includes var(ϵ). $L^2 = -2$ log-likelihood; χ^2 = the log-likelihood ratio chi-square statistic for the contrast between the current model and the previous model; *df* = the degrees of freedom associated with χ^2.
* $P \leq .001$.

distribution. Thus, we make this additional normality assumption in order to facilitate statistical inference.[20]

We estimate a series of parallel models for the T, S, and B dependent variables with the same sample restrictions associated with these dependent variables for models reported in tables 1 and 2. In table 3, we present goodness-of-fit statistics for three such models, which are modifications of equation (7). The first model is the variance components model in which all λ's, μ_1, μ_2, μ_4, and μ_5 are constrained to be zero. Its departure from the regionally homogeneous model (model 2 of table 1) is the introduction of a random component for the intercept (i.e., μ_0). Nine parameters, seven coefficients (α_0, α_1, α_2, α_3, α_4, α_5, and α_6) and two residual variances (var[ϵ] and var[μ_0]) are estimated for this model. For the random coefficients model, we further allow four μ's (μ_1, μ_2, μ_4, and μ_5) to contribute to their corresponding β coefficients. Since the two models are nested, we can assess the improvement in goodness of fit by

[20] Charles Manski pointed out to us that, asymptotically, there is no efficiency gain in reiterating the generalized least squares estimation of the variance-covariance matrix of the composite residual. However, for the convenience of statistical inference, we used the iterative procedure so that we could interpret IGLS estimates as asymptotic maximum-likelihood estimates. We used the computer program ML3 for estimation (for a review of estimation methods and computer programs for multilevel models, see Hox and Kreft [1994]).

taking the difference in the log-likelihood ratio statistic (L^2) to form a chi-square test. As shown in the last two columns, the contrast yields a chi-square test statistic with four degrees of freedom. For each of the three dependent variables, the random coefficients model is preferred to the simpler variance components model, since the chi-square test statistic is significant.

In the full model, we further allow our city-level variable z to affect individual-level coefficients. That is, λ_0, λ_1, λ_2, λ_4, and λ_5 are now allowed to be freely estimated. The full model consumes five more degrees of freedom than the random coefficients model. Again, we use a chi-square test (with $df = 5$) to assess the improvement in goodness of fit as we move from the random coefficients model to the full model. For all of the three dependent variables, the chi-square test is in favor of the full model at a P value less than or equal to 0.001. Note that the chi-square tests reported in table 3 are "composite" tests in the sense that they test several hypotheses simultaneously. Rejection of the simpler model thus does not mean that the additional parameters in the more complicated model should all be included. Although we initially tried to simplify the full model by trimming unnecessary parameters, we decided to report the full model to maintain consistency across the three dependent variables and discuss significance tests associated with individual parameters. For all three dependent variables, the estimated parameters of the full model and their estimated standard errors are reported in table 4.

We first examine the estimated parameters for the total earnings (T) equation. The exponential transformation of the estimated microlevel intercept (exp[6.384] = 592.3) should be interpreted as the mean earnings for a hypothetical group: male nonmembers with no schooling and no work experience living in a city that experienced no economic growth. In our approach of allowing the intercept to vary across cities, the intercept has a structural component ($\lambda_0 z$) and a random component (μ_0). The estimated λ_0 (under "Micro-macro interactive coefficients" in table 4) is significantly positive. This is to be expected, as more economic growth generates more economic wealth for redistribution to workers in the local labor force. Since z varies from 0.19 to 1.16 in our data, the estimated λ_0 of 0.685 would contribute between 0.130 and 0.795 to the baseline intercept term of 6.384. This result shows clearly that economic growth has a very significant effect on the *level* of earnings. The decompositional results from the second and third columns of table 4 show that the increase in the level of earnings associated with economic growth occurs mostly through the bonus share component rather than through regular salaries and wages (0.440 vs. 0.261, respectively). That is, the

TABLE 4

ESTIMATED PARAMETERS OF THE FULL MULTILEVEL MODEL FOR THREE DEPENDENT VARIABLES

	TOTAL EARNINGS (T)		SALARY/WAGE (S)		BONUS SHARE (B)	
	Parameter	SE	Parameter	SE	Parameter	SE
Microlevel coefficients:						
Intercept (α_0)	6.384	.059	5.937	.043	.443	.044
Years of schooling (α_1)	.029	.003	.030	.002	−.002	.002
Experience (α_2)	.045	.001	.037	.001	.007	.001
Experience² (α_3)	$(-6.35)10^{-4}$	$(2.30)10^{-5}$	$(-3.46)10^{-4}$	$(1.75)10^{-5}$	$(-2.71)10^{-4}$	$(1.75)10^{-5}$
Party member (1 = yes) (α_4)	.071	.019	.084	.014	−.013	.015
Gender (1 = female) (α_5)	−.332	.028	−.242	.023	−.065	.017
Gender × years of schooling (α_6)	.021	.002	.013	.001	.008	.001
Micro-macro interactive coefficients:						
Intercept (λ_0)	.685	.116	.261	.085	.440	.087
Years of schooling (λ_1)	−.017	.006	−.005	.005	−.009	.004
Experience (λ_2)	−.004	.002	−.005	.001	.000	.001
Party member (1 = yes) (λ_4)	.029	.039	−.014	.029	.040	.031
Gender (1 = female) (λ_5)	−.009	.043	.053	.035	−.087	.020
Macrolevel variance components:						
Intercept [var(μ_0)]	$(2.31)10^{-2}$	$(5.20)10^{-3}$	$(1.13)10^{-2}$	$(2.71)10^{-3}$	$(1.26)10^{-2}$	$(2.76)10^{-3}$
Years of schooling [var(μ_1)]	$(2.95)10^{-5}$	$(1.29)10^{-5}$	$(2.54)10^{-5}$	$(9.05)10^{-6}$	$(1.30)10^{-5}$	$(6.57)10^{-6}$
Experience [var(μ_2)]	$(2.47)10^{-6}$	$(1.20)10^{-6}$	$(1.83)10^{-6}$	$(7.77)10^{-7}$	$(1.28)10^{-6}$	$(6.62)10^{-7}$
Party member (1 = yes) [var(μ_4)]	$(8.92)10^{-4}$	$(6.14)10^{-4}$	$(3.52)10^{-4}$	$(3.15)10^{-4}$	$(7.67)10^{-4}$	$(4.00)10^{-4}$
Gender (1 = female) [var(μ_5)]	$(2.84)10^{-3}$	$(8.80)10^{-4}$	$(2.18)10^{-3}$	$(6.15)10^{-4}$	.000	.000
Microlevel variance component:						
var(ϵ)	.108	.001	.061	.001	.061	.001

NOTE.—For all of the three dependent variables, the full multilevel model in the form of eq. (7) is estimated. Goodness-of-fit statistics for these three models are reported in the "Full model" rows in table 3.

main reason for higher earnings in fast-growing cities is that workers living in such cities tend to receive a higher proportion of earnings in the form of bonuses and subsidies.

The effect of schooling on total earnings is 0.029 for males and 0.050 for females when economic growth is held at zero. Surprisingly, the education effect on total earnings is negatively related to economic growth (z), as the estimated λ_1 is -0.017 with an estimated standard error of 0.006. Within the observed range of variation, z contributes between -0.003 and -0.020 to the baseline education effect of 0.029. Although the negative influence of economic growth does not reverse the sign of the education effect, it can reduce it by as much as two-thirds. While λ_1 remains negative in both the S equation and the B equation, it is only statistically different from zero in the B equation. Thus, our first hypothesis is contradicted by the multilevel model for total earnings: economic growth is not positively related to the returns to education; instead, the relationship is negative. Furthermore, our decomposition results show that the negative impact of economic growth on the return to schooling works mainly through bonuses and subsidies. Not only does the bonus share not increase with education, it actually decreases with education in faster-growing cities.

Consistent with earlier results assuming regional homogeneity, work experience has a generally positive but concave effect on logged earnings. What is particularly interesting, however, is that the linear part of the experience effect is associated with economic growth in the same way as the education effect: λ_2 is estimated to be negative. In developing hypotheses earlier, we were not sure how to predict the relationship between the experience effect and economic growth. With the results reported in the first column of table 4, we are inclined to group education and experience together under a broad category of human capital. In the human capital framework, education and experience are the two most important dimensions of human capital, respectively representing formal and informal training (Mincer 1974). However, the similarity between experience and education is limited to the model for total earnings. Unlike the case of education, the interaction effect between experience and economic growth (λ_2) is significantly negative (-0.005) for the regular salary/wage education but insignificant for the bonus share equation. That is, work experience is a less important factor for determining a regular salary or wage in faster-growing cities than in slower-growing cities. This negative relationship suggests that government-sponsored seniority systems for setting salary and wage scales have been weakened in faster-growing cities.

We also observe from table 4 that the effects of party membership and gender on total earnings are unrelated to economic growth. For party membership, the absence of an effect holds true after decomposition. For gender, however, the story is a little more complicated. The gender gap in bonus share is exacerbated by economic growth. This tendency is somewhat offset by the narrowing effect of economic growth on the gender gap in regular salary and wages, although the latter effect is not statistically significant. Combined, these two opposing forces have no effect on total earnings.

REGIONAL VARIATION IN EARNINGS INEQUALITY

Our multilevel analysis properly models how earnings determination at the individual level varies across cities. We are still left with the question of how overall inequality varies across cities, and thus we are led back to the city as our unit of analysis. Our task in this section is to examine (1) whether overall inequality increases as a function of economic growth and (2) how regional variation in the earnings function affects overall inequality. For this purpose, we discuss two methodological strategies in some depth. The first is to devise appropriate city-level measures of inequality. The second is to draw inferences concerning city-level inequalities from individual-level earnings equations. Except where stated explicitly, in this section earnings refers to total earnings ($Y = Y_1 + Y_2 + Y_3$).

As Allison (1978, p. 867) remarked, "Perhaps the most commonly used measure of inequality is the Gini index." The Gini index can be defined as a measure of dispersion divided by twice the mean:

$$G = \left[1/n^2 \sum_{i=1}^{n} \sum_{j=1}^{n} |y_i - y_j| \right] \Big/ (2\mu_y), \tag{8}$$

where $\mu_y = E(y)$, and i and j refer to any two data points in the population. The Gini index is bounded between zero, which indicates absolute equality, and one, which indicates absolute inequality. For properties of the Gini index, the reader is referred to other sources (e.g., Allison 1978; Lerman and Yitzhaki 1984). Note that the definition of Gini presumes population data. In many research settings, including this one, we have access only to sample data. There are two solutions to this problem. The first solution is to ignore the difference between a population and a sample and apply the population formula to sample data. We call this solution "sample analog" (SA) estimation. For convenience, we follow the advice of Lerman and Yitzhaki (1984) in calculating SA estimates of G:

$$G_s = 2\,\mathrm{cov}[y, F(y)]/\bar{y}, \tag{9}$$

where $F(y)$ is the empirical cumulative distribution of y after y is rank ordered, and $\bar{y}$ is the sample mean of y. The second solution is to assume a parametric distribution underlying the dependent variable and derive a maximum-likelihood estimator. Our maximum-likelihood estimator is based on the assumption that the dependent measure y follows a lognormal distribution, as this assumption greatly simplifies Gini to a monotonic transformation of the standard deviation of logged y (Allison 1978, p. 874):

$$G_m = 2\Phi[S_{\log(y)}/(2^{1/2})] - 1, \tag{10}$$

where $S_{\log(y)}$ is the standard deviation of $\log(y)$, and $\Phi(\cdot)$ is the cumulative distribution function for a standard normal variable.

For each city, we calculate G_s and G_m according to equations (9) and (10) and list them in appendix table A1. With few exceptions, G_s and G_m vary between 0.17 and 0.28. For the whole data set, $G_s = 0.230$ and $G_m = 0.240$. By international standards, these figures are very low (Psacharopoulos 1981; Executive Yuan 1990, p. 19). However, we are not so much interested in the levels of Gini as in the variation of Gini across different cities in China. In particular, we want to examine the relationship between Gini or other measures of overall inequality and economic growth. Before we delve into this exercise, we would like to draw inferences a priori.

One dominant view of economic development and inequality is Kuznets's (1955) thesis that inequality follows an inverted U shape: inequality initially rises in the early and intermediate stages of development and then eventually declines with continued development. Strictly speaking, however, Kuznets's thesis is inapplicable to our study because Kuznets was concerned with total inequality rather than urban inequality. In fact, rural-to-urban migration is one of the key reasons behind Kuznets's conjecture that inequality increases during industrialization. However, Kuznets's thesis has been incorporated into a general modernization theory contending that increases in inequality accompany rapid economic growth in developing countries (Nee 1991, p. 277; Crenshaw and Ameen 1994, p. 2). While this conjecture is consistent with the experience of some countries such as Brazil (Fishlow 1972), Taiwan's recent history has clearly proven the opposite: Gini coefficients dropped gradually from 0.321 in 1964 to 0.277 in 1980 and then steadily climbed to 0.312 in 1990 (Executive Yuan 1990, p. 15). The example of Taiwan is relevant in three ways. First, while most studies confirming Kuznets's thesis are based on cross-sectional data (Gillis et al. 1987; Nee 1994, p. 7), high-quality trend data are available for Taiwan. Second, being culturally Chinese, Taiwan provides a natural reference for comparison with China. Third, Taiwan has had low levels of inequality comparable to China's. If Taiwan did not experience a rise in inequality during its early stages of rapid development during the 1960s and 1970s, there is reason to believe that inequality can be kept low during China's economic reforms.

Note that the two different methods of calculating Gini yield similar results, as the correlation between G_s and G_m is 0.932. Thus, the choice between the two appears inconsequential. For ease of computation and interpretation, we will focus our attention on G_m, which is in essence a monotonic transformation of the standard deviation of log(y). In fact, the standard deviation or the variance of log(y) can be used as a direct measure of inequality (Allison 1978). Labor economists (Chiswick 1971; Fishlow 1972; Lam and Levison 1992) have shown that total inequality can be decomposed based on the individual-level determinants of earnings. For ease of illustration, let us work with a simple case where there is no interaction effect between education and gender. That is, we set $\beta_6 = 0$. For each city we take the variance function on both sides of the equation and then take the partial derivatives of log Y with respect to β_1 and β_2:

$$\partial V(\log Y)/\partial\beta_1 = 2\beta_1 V(X_1) + 2\beta_2 \operatorname{cov}(X_1, X_2) + 2\beta_3 \operatorname{cov}(X_1, X_2^2) + 2\beta_4 \operatorname{cov}(X_1, X_4) + 2\beta_5 \operatorname{cov}(X_1, X_5); \quad (11a)$$

$$\partial V(\log Y)/\partial\beta_2 = 2\beta_2 V(X_2) + 2\beta_1 \operatorname{cov}(X_1, X_2) + 2\beta_3 \operatorname{cov}(X_2, X_2^2) + 2\beta_4 \operatorname{cov}(X_2, X_4) + 2\beta_5 \operatorname{cov}(X_2, X_5). \quad (11b)$$

We now make another simplifying assumption that the covariances do not vary across cities so that we can substitute estimates from the entire sample into the equation.[21] As an approximation, we also make use of our earlier parameter estimates under the assumption of regional homogeneity (i.e., model 1 of table 1). Thus, we reduce equation (11) into

$$\partial V(\log Y)/\partial\beta_1 = -.02799 + 19.4992\beta_1; \quad (12a)$$

$$\partial V(\log Y)/\partial\beta_2 = -6.0645 + 210.1464\beta_2. \quad (12b)$$

That is to say, as long as β_1 is above a small threshold of 0.0144, which is generally true for our data, earnings inequality is a positive and increasingly positive function of the rate of return to education. Likewise, as long as β_2 is above a small threshold of 0.0289, which is also true, earnings inequality is a positive and increasingly positive function of the return to the linear component of work experience.

Hence, we infer that macrolevel factors that increase the rate of returns to education and work experience also increase overall inequality. At low rates of return, increases in overall inequality due to increases in returns to human capital are slow. At high rates of return, however, increases in overall inequality due to increases in returns to human capital are very

[21] Violation of this assumption would mean that there are three-way interactions among city and a pair of X variables.

fast, as the relationship between the two is nonlinear and convex. It is for this reason that we stated that Nee's market transition theory implies an increasing trend in overall inequality (see n. 4).

In table 5, we present the correlation matrix for various city-level indices and regression parameters. As expected, faster economic growth is correlated with higher mean earnings (correlation = 0.577) and higher mean logged earnings (correlation = 0.502). Furthermore, the correlation between economic growth and the mean of our bonus share measure is much higher (0.522) than that between economic growth and the mean of logged salary and wages (0.216). Note that the city-level measure of bonus rate (reported in app. table A1) is highly correlated with the mean of our measure of the bonus share (0.980), an indication that B captures intercity variation in the importance of bonuses and subsidies relative to regular salary and wages.

We note the negative correlation (−.407) between R^2 from the city-specific baseline model and economic growth (z). This negative correlation suggests that earnings determination varies less systematically and conforms less to the human capital model in cities where economic growth has been faster than in cities where economic growth has been slower. Much of this negative correlation is due to bonuses and subsidies, as R^2 has relatively large negative correlations with the mean of B and the city-level bonus rate (−0.515 and −0.503, respectively) but a small positive correlation with logged salary/wage (0.103).

Consistent with results from the multilevel analysis, faster economic growth is correlated with a lower return to education (correlation = −0.337) and with a lower return to work experience (correlation = −0.115).[22] Of special interest in table 5 are the correlations between economic growth and several measures of inequality: economic growth has low correlations with the SA and ML estimates of the Gini index (0.287 and 0.236, respectively). These results do not lend strong support to the modernization view that high inequality is a necessary price to pay for rapid economic growth. In brief, cities with faster-growing economies are characterized by higher mean earnings (particularly higher bonuses and subsidies), lower returns to education and experience, less conformity to the baseline model for earnings determination, and slightly higher overall inequality.

Given equation (12), these results imply that education and experience work to *depress* the overall inequality in faster-growing cities. That is, the correlation between economic growth and overall inequality would

[22] We constrain the effect of education to be parallel between men and women and the effect of experience to be linear in order to examine simple relationships between these effects and economic growth.

TABLE 5

CORRELATIONS AMONG CONTEXTUAL VARIABLES AND ESTIMATED STATISTICS ACROSS 55 CITIES IN CHINA

	z	log p	$\bar{y}$	$\overline{\log y}$	$\bar{S}$	$\bar{B}$	BR	G_m	G_s	R^2	G_r	β_1	β_2
Economic growth (z)	1.000												
Logged population size (log p)	−.151	1.000											
Mean of earnings ($\bar{y}$)	.577	.060	1.000										
Mean of logged earnings ($\overline{\log y}$)	.502	.077	.967	1.000									
Mean of logged salary/wage ($\bar{S}$)	.216	.020	.728	.734	1.000								
Mean of logged bonus share ($\bar{B}$)	.522	.081	.661	.699	.031	1.000							
Bonus rate (BR)	.517	.089	.634	.673	.018	.980	1.000						
Gini, ML estimate (G_m)	.287	.047	.178	.025	−.099	.182	.237	1.000					
Gini, SA estimate (G_s)	.236	.029	.139	−.023	−.144	.145	.201	.932	1.000				
R^2 (R^2)	−.407	−.060	−.341	−.290	.103	−.515	−.503	−.256	−.293	1.000			
Residual Gini (G_r)	.386	.047	.254	.103	−.129	.321	.368	.959	.906	−.512	1.000		
Coeff. of education (β_1)	−.337	−.188	−.185	−.225	−.040	−.276	−.263	.206	.167	.451	.041	1.000	
Coeff. of experience (β_2)	−.115	.077	−.134	−.166	−.013	−.200	−.145	.395	.294	.358	.241	.445	1.000

NOTE.—ML = maximum-likelihood estimation and SA = sample-analog estimation. Regression coefficients (β's) refer to the least-squares estimates of eq. (1) for each of the 55 cities, with the constraint that $\beta_3 = \beta_6 = 0$. See app. table A1 for an explanation of bonus rates. City-level bonus rates are reported in app. table A1. R^2 and G_r come from app. table A2.

be higher if the returns to human capital were constant across cities or positively affected by economic growth. One piece of evidence that confirms this reasoning is the higher correlation (0.386) between economic growth and the residual Gini after controlling for the effects of the independent variables through a regression analysis.[23]

DISCUSSION

To recapitulate, our results are negative: We failed to find higher returns to education and lower returns to party membership in cities that experienced faster economic growth. Instead, for total earnings the returns to education as well as to work experience are negatively associated with economic growth, and the effects of party membership and gender are regionally invariant. Furthermore, we find the correlation between economic growth and overall earnings inequality to be moderate, in part due to the equalizing influence of the negative relationship between the returns to human capital factors and economic growth.

These new results contradict expectations derived from the literature and thus present an interesting empirical puzzle. Very often, it is new empirical findings that serve as the ultimate arbitrator of theoretical disputes and set the course for future theoretical work. Although our primary interest in this paper is to report surprising empirical results, we would like to offer our interpretation in this section.

While data quality can always be suspect, we believe that it is an unlikely source for the negative findings.[24] Compared to data used in similar studies on earnings inequality in China, CHIP is superior for two reasons. First, CHIP is a large national sample; in contrast, most earlier studies (with the few exceptions of Khan et al. 1992; Griffin and Zhao 1993; Nee 1994, 1996) are based on regional samples. Second, CHIP used an unusually detailed survey instrument to collect information pertaining to earnings, including all forms of subsidies (see Griffin and Zhao 1993). In addition, our results are robust in the sense that our qualitative conclusions do not depend on our multilevel model. As shown in table 5, the returns to education and experience estimated via OLS separately for each city (with β_3 and β_6 set to zero) are negatively correlated with the measure of economic growth. Although we concede the possibility that economic growth may have been

[23] The residual Gini measures the amount of inequality unexplained by the city-level baseline model. It is calculated by applying eq. (10) to the root mean square error (RMSE), the estimated standard deviation of unexplained log Y.

[24] For example, while we suspect that a few obvious outliers in Taiyuan and some other cities are probably due to misreporting, deleting these outliers does not alter our substantive conclusions.

measured with error, there is no reason to believe that measurement errors are so systematic as to bias our results. We have confirmed the validity of our measure of economic growth both by a city-by-city check for face validity and by its high correlation with mean earnings (table 5). One likely source of measurement error for economic growth is possible changes in geographic boundaries between 1985 and 1988 (see n. 18). Such changes should favor smaller cities, which have expanded more rapidly in recent years. That is, it is possible that our measure of economic growth for small cities is inflated due to the problem of boundary reclassification between 1985 and 1988. Table 5 shows, however, that the logarithm of population size in 1988 has only a weak negative correlation with the measure of economic growth (-0.151). We should not attribute all of this correlation to boundary reclassification, since we know that smaller cities have actually experienced faster rates of real growth since the onset of the reforms. In any event, it is safe to conclude that this correlation is too small for city size or measurement errors associated with city size to explain away the role of economic growth.

One likely explanation for the negative findings is that Nee's market transition theory and Chiswick's economic reasoning may be applicable to rural China but inapplicable to urban China. As observed by Róna-Tas (1994, p. 44), Nee's support for market transition theory has so far been restricted to data from rural China; whether conclusions drawn from studies of rural workers can be generalized to urban workers remains an open question. Vast differences exist between the urban and rural sectors of China, in large part because the two sectors have been governed by distinct administrative policies with tight restriction on rural-to-urban migration. The scholarship on contemporary China has nearly always made a sharp distinction between the rural and urban sectors, treating them as separate entities. In a recent study of wages that combines the two, Peng (1992) finds that the Chinese wage system is characterized far more by the rural-urban distinction than by the public-private distinction, as wage determination in the rural public sector is similar to that in the rural private sector but quite different from that in the urban state sector. One prominent rural-urban difference is that, prior to economic reforms, Chinese agriculture was collectivized into communes but was never nationalized, whereas the state owned virtually all of the important enterprises in urban China. Indeed, one may make a case that industrial reforms in China are more partial—that is, farther away from the ideal of "markets"—than agricultural reforms.

It seems plausible that our negative findings can be attributed to "partial reforms" in urban China, as the transition to markets is far from complete. This argument, however, is untenable for three reasons. First, as Róna-Tas (1994, p. 44) argues, it is too easy to attribute negative findings to "partial reforms" and thus to make market transition theory unfalsifiable. Second, if the negative findings are indeed attributable to partial reforms in urban China, we are interested in knowing how partial reforms have brought about a regional variation in earnings determination that contradicts our theoretical expectations. Third, Nee (in this issue) reports that for rural China the effect of education on income is significantly positive for inland provinces but nil or negative for coastal provinces, a finding that he found unexpected (p. 940) but that is consistent with our results.

We attribute our negative findings to the lack of a true labor market in 1988 urban China. Restated, the relationship between employers and employees remains essentially in the old socialist order, although economic reforms have generated new manifestations of this relationship. As pointed out by Lin and Bian (1991, p. 661) in a study of status attainment in urban China, "status attainment in China is more geared to work units than to occupations per se." That is to say, *where* one works matters a great deal and probably more than *what* tasks one performs in the workplace. Similar to occupation in western societies, one's work unit affiliation in China reveals much about one's social status. Before economic reforms, inequalities generated by work units were mainly manifested in terms of housing, access to facilities (such as bathhouses and childcare centers), the ability to confer or seek favors (*guan-xi* or "connections"), and other fringe benefits. Since the reforms, however, the benefits that workers derive from work units have become more monetary. Walder's (1986, 1987, 1992*a*, 1992*b*) work has clearly demonstrated that work organizations commonly engage in handing out monetary benefits to their workers, sometimes over the upper limits set by the government. When work organizations have more resources at their disposal, these resources are directly translated into financial benefits to workers in the form of bonuses and subsidies, regardless of workers' actual contributions. That is to say, intraorganizational inequality is kept low while interorganizational inequality may have risen, primarily through the distribution of bonuses and subsidies.

Let us illustrate the situation with a hypothetical example. Say two incumbents ("a" and "b") of the same occupation have identical qualifications and are thus substitutable for each other. The actual amount they are paid may vary greatly depending on the abilities of their work units (A and B, respectively) to afford bonuses and subsidies. It is possible that organization A pays worker a 1,000 yuan (about 53% of average urban earnings in 1988) more than organization B pays worker b simply because A has more money at its disposal than does B. In fact, A pays all of its workers about 1,000 yuan above the level B pays its workers. Because there is no labor market in urban China, workers working for organization B cannot quit to move to organization A in order to get higher pay; neither can organization A replace its workers with those working for B in order to lower its labor costs. In fact, as shown later, organization A has no incentives to lower its labor costs. Hence, pay inequities across work organizations will not disappear or even approach equilibrium. Although greatly simplified, this scenario captures much of the reality described by recent scholarship, particularly Walder's, on social inequalities and work organizations.

How do we account for organization A's seemingly irrational practice of paying its workers more than "necessary" in the form of bonuses and subsidies and thus foregoing higher profits? Walder (1992*a,* p. 526) persuasively eliminates as an explanation the need to use incentives for retaining skilled workers, as the labor market is essentially nonexistent and job mobility is extremely low in urban China (Walder 1986; Davis 1992). There are two explanations, one economic and the other sociological. The economic explanation is based on the fact that socialist economies typically operate under what Kornai (1986, 1989) characterizes as a "soft budget" due to vague definitions of property rights. As Kornai's work demonstrates, under a soft budget, it is quite rational for work organizations to spend as much as possible since they are constantly bargaining to obtain or protect economic resources from the central government. Rewarding workers with extra pay or other tangible benefits increases the cost of production and thus reduces the amount of profit being sucked away into the government's vast redistributive channel.

The sociological explanation argues that social structures in socialist societies are tightly organized around work organizations (Walder 1986). In socialist economies, employment is not a market relationship in terms of payment for labor or services performed. Rather, "employment plays a welfare role," with work organizations being responsible for the overall well-being of their employees (Walder 1986, p. 11). Although the government has attempted to change this employer-employee relationship in recent years (e.g., housing has been commercialized in limited areas), the norm of employer paternalism is so entrenched in Chinese society that the attempt has not been successful. In fact, workers' dependency on work organizations has not decreased but increased, since bonuses and subsidies, which are set by the work units, have constituted an increasingly large fraction of earnings over the years (Walder 1992*b*). In our CHIP data, we have found bonuses and subsidies to comprise higher proportions of total employment earnings in faster-growing cities than in slower-growing cities.

Although all work organizations have an incentive to "overpay" their workers, not all of them are in a position to do so. As a consequence of economic reforms, organizations are allowed to reward their workers only when they can find the financial resources needed for doing so. However, this pay inequality generated by work organizations appears to critics as "irrational" or "unfair" because profit differences between organizations can be arbitrary. A Chinese observer (Zhao 1994, p. 113) points out that profit levels are affected by many "external conditions" above and beyond workers' productivity. Such external conditions include the dual-pricing system, domestic and international markets, administrative mismanagement, and the government's financial policies toward certain districts, certain industries, and certain enterprises. From this, Zhao concludes, "Excessive earnings inequalities across districts, industries, and work units attributable not to differences in worker productivity but to these [external] factors have generated the masses' discontent. On the surface, this discontent reflects the unfairness of distribution; in actuality it reflects the unfairness of process. From this, we think that the government . . . should provide all workers an equal opportunity for fair competition" (p. 113, our translation from Chinese). By "the unfairness of process" Zhao means that the affiliation with highly profitable enterprises, which is the major determinant of earnings, is determined mainly through administrative channels and not through market forces.

If interorganizational inequalities are largely due to external factors that are related to economic disequilibrium, intraorganizational inequalities can be even more arbitrary. Walder (1987, 1992*b*) and Shirk (1989) both show that workers prefer equal distribution of pay and reject any notion of linking pay and productivity within organizations. Let us recall that earnings in 1988 urban China essentially consist of two parts roughly equal in importance: (*a*) regular salary or wages and (*b*) bonuses and subsidies. While the first part is set by the government and varies little, the second part is set by individual work units. Note that the returns to education and experience are explicitly recognized in the government's pay schedule. Walder's and Shirk's work suggests that human capital factors do not matter much in the determination of bonuses and subsidies within an organization. If this is the case, the more capable organizations are of giving monetary benefits to workers, the higher the intercept of the earnings equation but the *flatter* the slopes of human capital factors.[25] Large geographic variations in economic growth and thus financial capability explain why the returns to education and experience are lower and average earnings are higher in cities where economic growth has been faster.

Direct evidence has yet to be collected that would show workers' high dependency on the financial conditions of their work units for bonuses

[25] This is so because our dependent variable in the earnings equations is in the logarithm scale. A fixed amount of extra pay translates into a smaller proportional increase in earnings for the more educated than for the less educated.

TABLE 6

SIGNIFICANCE OF VARIOUS FACTORS IN DETERMINING REGULAR SALARY AND BONUS SHARE: RESULTS FROM AN ANALYSIS OF VARIANCE

		Salary (*S*)		Bonus Share (*B*)	
Source	*df*	*SS*	*MS*	*SS*	*MS*
By explanatory variable:					
City	54	142.258	2.634	251.588	4.659
Education	7	43.905	6.272	.778	.111
Experience	45	429.793	9.551	29.213	.649
Party member	1	5.402	5.402	.093	.093
Gender	1	9.145	9.145	.129	.129
Gender-education interaction	7	5.379	.768	.810	.116
Employer ownership	8	20.301	2.538	4.199	.525
Occupation	7	7.776	1.111	.810	.116
Industry	16	11.200	.700	5.989	.374
Employment term	4	4.750	1.188	7.279	1.820
Model	150	1,124.339	7.496	326.907	2.179
Error	15,437	904.494	.059	927.312	.060
Total	15,587	2,028.833		1,254.219	
R^2 (%)		55.42		26.06	

NOTE.—N = 15,588. All explanatory variables are entered into the analysis of variance as categorical variables. *df* stands for degrees of freedom, *SS* sum of squares, and *MS* mean of squares (*MS* = *SS*/*df*). For explanatory variables, *SS* refers to the partial sum of squares with other variables controlled for. Missing values are coded into a separate category for employer ownership, occupation, industry, and employment term.

and subsidies. The CHIP data, however, reveal the predominance of region, compared to other factors, in determining bonuses and subsidies. In table 6, we present analysis of variance (ANOVA) results with the logged regular salary/wage (*S*) and the bonus share (*B*) as dependent variables. The objective is to see the relative importance of various factors. For this analysis, we include a few variables not considered previously: employer ownership, occupation, industry, and employment term.[26] As a conservative strategy, all explanatory variables are entered as categorical variables. The main entries of table 6 are partial sums of squares and partial means of squares associated with each independent variable. It is clear that patterns determining the two dependent variables are drastically different. Let us first use the mean of squares as a crite-

[26] The variable "employment term" is designed to measure the nature of the employment relationship with four categories: "permanent," "temporary," "most of time spent working on private enterprises," and "other." The survey instruments measuring industry and occupation were very crude. For all of the variables, original codes are preserved and entered into the analysis. Missing values are coded into a separate category for employer ownership, occupation, industry, and employment term.

rion, which adjusts for degrees of freedom. For the logged regular salary/wage (S), city is the sixth most important factor after (in order of importance) experience, gender, education, party membership, and employer ownership. For the bonus share (B), city is by far the most important factor: a 4.659 mean of squares is attributable to intercity differences, with the next highest mean of squares being 1.820 (for employment term). Similar to results shown in table 2, the bonus share is less "explainable" than regular salary/wage. For the B model, R^2 is 26.06%; R^2 is 55.42% for the S model. To the limited extent that it is explainable, the bonus share is mostly determined by geography: at least 77.0% of the small R^2 for the B model is attributable to intercity variation.[27] For S, the comparable figure is merely 12.7%.

These ANOVA results in table 6 are consistent with our interpretation that bonuses and subsidies are distributed to workers mainly for their affiliations with profitable work units. The share of bonuses and subsidies among total employment earnings does not reward human capital factors because bonuses and subsidies are regulated neither by the old socialist regime nor by a market regime. In the socialist regime, salary is set administratively as an increasing function of education and experience; in the market regime often regarded as the ultimate destination of Chinese economic reforms, employers would need to raise bids in order to compete for productive workers. Given the absence of true labor markets in 1988 urban China, moderate returns to human capital factors for the regular salary or wage suggest that salaries and wages are governed by the old socialist regime. In contrast, bonuses and subsidies are largely unregulated according to any clearly articulated logical system.

CONCLUSIONS

This study provides new and forceful evidence against certain conjectures currently popular in the literature. Capitalizing on large regional variations in the pace of economic reforms and consequently in economic growth, we have shown that the rates of return to education do not increase with economic growth, as Nee's theory of market transition and Chiswick's economic analysis predict. Instead, returns to education as well as to work experience are lower in cities with faster economic growth than in cities where growth has been slower. Nee's prediction that the significance of positional power declines with the progress of economic

[27] We use the term "at least" because the sums of squares in table 6 are partial sums of squares excluding joint contributions with other variables to the sum of squares of the model. That is why the sum of all partial sums of squares is less than the sum of squares of the model.

reforms is also unsupported, as we find the return to party membership to be invariant.[28] Our results are consistent with Róna-Tas's (1994) recent work, which clearly demonstrates the importance of political capital for maintaining an advantaged position during Hungary's transition to a market economy. In addition, our results show that the gender gap in total earnings is also unrelated to the pace of economic growth. The level of overall earnings inequality remains low in China and only slightly correlated with economic growth, in part because the tendency toward higher levels of inequality in faster-growing cities is offset by the lower returns to human capital variables in these cities.

Given the high quality of the data and the appropriateness of the statistical strategies used, we conclude that our negative results serve to reject the applicability of Nee's market transition theory, at least to urban China in 1988. Although one cannot really rule out the possibility that market transition theory will become more applicable with time, we do not resort to "partial reforms" as the default explanation, as it is tautological and unfalsifiable. Rather, we find our results to be congruent with economic and sociological explanations that give primary attention to institutions as agents of stratification in socialist and reforming-socialist economies.

In conclusion, we find that in terms of earnings determination, reform-era urban China cannot be simply characterized as falling along a continuum between a redistributive economy and a market or capitalist economy. Instead, we argue that the postreform Chinese economy has qualitatively unique features and should be studied accordingly. One unique feature is that, while there are free or semifree markets for goods and services, labor markets are virtually absent. A related feature is the vagueness in defining property rights and articulating the rules for profit sharing for nominally public enterprises. Given the continued acceptance of the ethos that work organizations should provide welfare to workers, profitability differentials are translated into large differentials in bonuses and subsidies across work organizations and thus regions. Variations in bonuses and subsidies cannot be adequately explained by a traditional human capital model. Ironically, the extra earnings in the form of bonuses and subsidies that epitomize economic reforms are even less subject to market forces than are regular salaries and wages. We view this irony as an unintended consequence of economic reforms in urban China, which until now have failed to effectively address the fundamental issues of property rights and labor markets.

[28] We would like to emphasize again that party membership is not a direct measure of positional power but a proxy for positional power.

APPENDIX

TABLE A1

SUMMARY MEASURES OF EARNINGS INEQUALITY AND ECONOMIC INDICATORS BY PROVINCE AND CITY

	1988 CHIP Micro Data								1985–88 CUS	
		Annual Earnings (Yuan)		Logged Earnings		Gini Index			Economic Growth	
	N	Mean	SD	Mean	SD	G_m	G_s	Bonus Rate	z	Annual Rate (%)
Total	15,862	1,871	1,077	7.439	.431	.240	.230	.45		
Beijing:										
Beijing	815	1,952	715	7.507	.388	.216	.196	.49	.24	8.3
Shanxi:										
Taiyuan	382	1,830	2,207	7.354	.477	.264	.281	.42	.26	9.1
Datong	336	1,645	828	7.318	.422	.234	.219	.40	.28	9.8
Yangquan	207	1,672	1,345	7.299	.449	.249	.260	.37	.19	6.5
Changzhi	200	1,587	582	7.293	.426	.237	.196	.41	.51	18.5
Jincheng	198	1,699	804	7.336	.453	.251	.252	.43	.79	30.1
Yuncheng	179	1,642	884	7.306	.433	.241	.231	.40	.57	20.9
Liaoning:										
Shenyang	619	1,815	642	7.454	.307	.172	.173	.45	.32	11.3
Dalian	613	1,887	555	7.504	.279	.156	.152	.47	.34	12.0
Dandong	182	1,780	618	7.438	.298	.167	.161	.43	.31	10.9
Jinzhou	371	1,631	521	7.348	.316	.177	.176	.40	.27	9.4

TABLE A1 (*Continued*)

		1988 CHIP Micro Data							1985–88 CUS	
		Annual Earnings (Yuan)		Logged Earnings		Gini Index			Economic Growth	
	N	Mean	SD	Mean	SD	G_m	G_s	Bonus Rate	z	Annual Rate (%)
Jiangsu:										
Nanjing	524	1,730	520	7.410	.312	.174	.161	.44	.43	15.4
Wuxi	443	1,865	553	7.485	.314	.176	.163	.49	.54	19.7
Xuzhou	340	1,692	554	7.380	.339	.189	.180	.40	.34	12.0
Changzhou	176	1,878	723	7.475	.359	.200	.186	.47	.50	18.1
Nantong	194	1,931	636	7.516	.324	.181	.162	.47	.52	18.9
Huaiyin	107	1,282	474	7.089	.372	.207	.209	.40	.56	20.5
Yancheng	90	1,765	526	7.423	.354	.197	.164	.45	.57	20.9
Yangzhou	290	1,843	1,946	7.420	.365	.204	.217	.45	.89	34.5
Anhui:										
Hefei	409	1,786	983	7.406	.401	.223	.197	.45	.48	17.4
Wuhu	173	1,712	788	7.381	.348	.194	.179	.43	.37	13.1
Bengbu	186	1,730	753	7.361	.452	.251	.237	.48	.35	12.4
Huainan	377	1,677	981	7.304	.494	.273	.258	.43	.32	11.3
Anqing	292	1,803	1,526	7.365	.465	.258	.267	.46	.35	12.4
Fuyang	91	1,172	510	6.983	.413	.230	.226	.40	.48	17.4
Tunxi	85	1,584	688	7.286	.416	.231	.212	.46	.38	13.5
Henan:										
Zhengzhou	384	1,613	638	7.324	.350	.196	.190	.35	.29	10.1
Kaifeng	194	1,393	473	7.183	.340	.190	.179	.36	.25	8.7
Luoyang	362	1,566	592	7.272	.437	.242	.214	.27	.43	15.4
Pingdingshan	182	1,500	560	7.246	.372	.207	.201	.34	.46	16.6
Anyang	285	1,430	661	7.165	.479	.265	.224	.33	.37	13.1
Xinxiang	269	1,520	585	7.247	.418	.233	.212	.37	.60	22.1
Zhumadian	124	1,537	586	7.273	.362	.202	.191	.39	.57	20.9
Nanyang	105	1,563	485	7.304	.328	.183	.168	.41	.39	13.9

Hubei:										
Wuhan	653	1,762	620	7.419	.343	.192	.174	.41	.28	9.8
Huangshi	196	1,798	537	7.440	.358	.200	.166	.45	.30	10.5
Shiyan	194	1,889	598	7.499	.302	.169	.159	.44	.41	14.6
Shashi	207	1,660	554	7.370	.296	.166	.160	.38	.37	13.1
Xiaogan	192	1,773	447	7.447	.268	.150	.138	.44	.75	28.4
Guangdong:										
Guangzhou	557	2,506	1,745	7.710	.454	.252	.250	.53	.41	14.6
Shaoguan	189	2,036	970	7.497	.558	.307	.239	.53	.42	15.0
Shenzhen	206	4,066	1,541	8.237	.410	.228	.196	.47	1.16	47.2
Shantou	97	2,036	1,515	7.437	.617	.338	.313	.52	.97	38.2
Foshan	309	3,282	2,026	7.995	.423	.235	.238	.68	.97	38.2
Zhanjiang	196	2,223	1,462	7.567	.536	.295	.270	.58	.75	28.4
Huizhou	220	2,318	1,017	7.655	.454	.252	.225	.56	1.04	41.4
Zhaoqing	207	2,229	1,180	7.592	.518	.286	.244	.58	.79	30.1
Yunnan:										
Kunming	575	2,024	693	7.558	.343	.192	.174	.50	.32	11.3
Dongchuan	151	1,833	1,310	7.412	.414	.230	.235	.34	.20	6.9
Qujing	80	2,023	581	7.569	.304	.170	.162	.48	.34	12.0
Yuxi	186	2,142	928	7.588	.400	.223	.223	.54	.61	22.5
Gejiu	192	1,834	1,040	7.431	.394	.220	.206	.37	.40	14.3
Dali	184	1,800	454	7.461	.272	.153	.142	.42	.39	13.9
Baoshan	191	1,713	482	7.409	.275	.154	.150	.43	.27	9.4
Gansu:										
Lanzhou	1,096	1,865	1,286	7.409	.497	.275	.251	.38	.22	7.6

NOTE.—N denotes the sample size, and G_m and G_s, respectively, denote the maximum-likelihood and sample-analog estimates for the Gini index of inequality. Bonus rate is defined (for those respondents with positive salary/wage) as $\overline{Y}_2/(\overline{Y}_1 + \overline{Y}_2)$, where $\overline{Y}_1$ denotes the average regular salary/wage and $\overline{Y}_2$ denotes the average cash bonuses/subsidies. Economic growth is measured by gross product value of industry (GPVI); $z = \log(\mathrm{GPVI}_{1988}/\mathrm{GPVI}_{1985})$, and the last column labelled "annual rate" converts z into annualized rate of growth.

TABLE A2

ESTIMATES OF THE MODIFIED HUMAN CAPITAL MODEL BY PROVINCE AND CITY

	β_0	β_1	β_2	β_3 ($\times 10^{-4}$)	β_4	β_5	β_6	R^2 (%)	RMSE	G_r
Beijing:										
Beijing	6.835	.018	.043	−6.485	.037	−.184	.005	28.7	.329	.184
Shanxi:										
Taiyuan	6.461	.023	.051	−8.198	.225	−.420	.032	40.5	.370	.207
Datong	6.748	.010	.054	−9.350	.075	−.893	.060	40.7	.328	.183
Yangquan	6.349	.036	.060	−9.651	−.090	−.472	.030	30.1	.381	.212
Changzhi	6.446	.019	.056	−7.355	.040	−.722	.058	35.0	.349	.195
Jincheng	6.860	−.008	.067	−12.22	.121	−1.090	.077	44.8	.342	.191
Yuncheng	6.697	.008	.040	−5.340	.030	−.564	.042	26.9	.377	.210
Liaoning:										
Shenyang	6.747	.024	.033	−4.142	.077	−.220	.015	34.3	.250	.140
Dalian	6.857	.019	.034	−4.276	.059	−.191	.013	35.6	.225	.126
Dandong	6.717	.038	.021	−1.232	.037	.000	−.006	32.8	.249	.140
Jinzhou	6.623	.023	.037	−4.746	.102	−.372	.024	46.3	.233	.131
Jiangsu:										
Nanjing	6.688	.017	.047	−7.346	.095	−.227	.013	48.4	.225	.126
Wuxi	6.716	.029	.034	−3.954	.067	−.070	.004	38.6	.248	.139
Xuzhou	6.794	.014	.044	−6.901	.091	−.557	.038	44.7	.254	.143
Changzhou	6.969	.003	.045	−7.132	.051	−.396	.025	31.4	.302	.169
Nantong	7.100	.004	.031	−4.214	.118	−.305	.018	34.4	.266	.149
Huaiyin	6.313	.036	.046	−7.419	.107	−.220	.017	30.1	.320	.179
Yancheng	6.769	.015	.041	−5.935	.101	−.199	.010	41.2	.281	.157
Yangzhou	6.821	.018	.032	−3.542	.041	−.284	.018	29.3	.310	.174
Anhui:										
Hefei	6.496	.036	.038	−4.608	.069	−.138	.009	34.9	.326	.182
Wuhu	6.611	.029	.035	−4.191	.026	−.073	.002	29.4	.298	.167
Bengbu	6.344	.040	.054	−7.099	−.010	−.275	.010	41.3	.352	.196
Huainan	6.871	.015	.036	−4.849	.068	−.940	.061	38.3	.391	.218
Anqing	6.558	.024	.049	−6.970	.038	−.281	.021	26.8	.402	.224
Fuyang	6.679	.013	.021	−2.211	.111	−.429	.021	21.5	.379	.211
Tunxi	6.662	.019	.031	−2.462	.105	−.935	.090	45.2	.319	.179

Henan:										
Zhengzhou	6.848	.017	.021	−2.049	.116	−.433	.027	35.9	.282	.158
Kaifeng	6.441	.031	.033	−4.345	.194	−.015	−.001	28.3	.293	.164
Luoyang	6.311	.021	.065	−11.01	.185	−.463	.034	41.2	.338	.189
Pingdingshan	6.765	.013	.027	−2.902	.172	−.421	.024	36.1	.302	.169
Anyang	6.322	.023	.056	−8.579	.075	−.269	.009	39.0	.378	.211
Xinxiang	6.487	.019	.041	−5.194	.146	−.362	.028	31.9	.349	.195
Zhumadian	6.506	.027	.047	−7.853	.092	−.320	.017	36.1	.297	.166
Nanyang	6.599	.030	.030	−3.381	.055	−.165	.002	38.5	.265	.148
Hubei:										
Wuhan	6.546	.030	.042	−6.195	.072	−.243	.016	38.0	.272	.152
Huangshi	6.465	.014	.071	−11.89	.020	−.402	.036	48.2	.262	.147
Shiyan	6.867	.016	.032	−4.206	.084	−.217	.018	22.4	.270	.151
Shashi	6.781	.017	.028	−2.937	.042	−.057	.005	29.8	.252	.141
Xiaogan	6.672	.022	.044	−6.393	−.024	−.040	−.001	30.3	.228	.128
Guangdong:										
Guangzhou	7.271	.010	.029	−4.032	.132	−.335	.020	14.7	.421	.234
Shaoguan	6.498	.022	.078	−14.32	.159	−.446	.030	27.1	.484	.268
Shenzhen	7.168	.025	.082	−16.99	.070	−.226	.013	32.4	.342	.191
Shantou	6.268	.047	.065	−12.17	.034	1.007	−.104	22.7	.561	.308
Foshan	7.782	.002	.017	−1.804	.141	−.350	.020	14.8	.394	.220
Zhanjiang	6.654	.025	.060	−9.654	.085	−.079	−.004	22.1	.480	.266
Huizhou	6.855	.036	.034	−5.090	.153	−.217	.009	28.7	.389	.217
Zhaoqing	6.637	.034	.059	−9.176	−.002	−.370	.018	28.9	.444	.246
Yunnan:										
Kunming	6.619	.034	.046	−7.089	.060	.040	−.007	29.4	.290	.163
Dongchuan	6.544	.030	.036	−3.928	.175	−.374	.024	43.9	.317	.177
Qujing	7.129	.025	.011	−.008	.097	−.371	.029	40.8	.243	.137
Yuxi	7.327	−.003	.023	−2.750	.122	−.337	.021	18.5	.367	.205
Gejiu	6.320	.041	.062	−9.775	.015	−.237	.010	45.9	.295	.165
Dali	6.906	.026	.023	−2.944	.053	−.184	.003	31.2	.230	.129
Baoshan	6.517	.037	.041	−5.778	.009	−.054	−.005	50.4	.196	.110
Gansu:										
Lanzhou	6.413	.032	.056	−8.144	.120	−.679	.049	41.3	.382	.213

NOTE.—All β's are ordinary least squares estimates of eq. (1). RMSE stands for root mean squared error, and G_r for the maximum-likelihood estimate of the Gini index based on RMSE.

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