ELSEVIER

Available online at www.sciencedirect.com

SCIENCE DIRECT®

Journal of Comparative Economics 33 (2005) 644–663

Journal of COMPARATIVE ECONOMICS

www.elsevier.com/locate/jce

Has China crossed the river? The evolution of wage structure in urban China during reform and retrenchment

Simon Appleton [a,*], Lina Song [b], Qingjie Xia [c]

[a] *School of Economics, University of Nottingham, University Park, Nottingham NG7 2RD, UK*
[b] *School of Sociology and Social Policy, University of Nottingham, NG7 2RD, UK*
[c] *School of Economics, Peking University, Beijing, China 100871*

Received 10 August 2005

Available online 30 September 2005

Appleton, Simon, Song, Lina, and Xia, Qingjie—Has China crossed the river? The evolution of wage structure in urban China during reform and retrenchment

In this paper, we examine the determinants of urban wages in China from 1988 to 2002. We find increased returns to education but a decrease in the returns to experience. The 2002 data imply that the widening pure gender gap and the growth in the premium to Communist Party membership may have come to an end. The reform of the state-owned enterprise (SOE) sector and the shift in industrial structure out of heavy industry is shown to impact wages of workers within those sectors. We use recall panel data for 1998 to 2002 to provide fixed effects estimates of the impact of sector ownership, Communist Party membership and unemployment on wages. *Journal of Comparative Economics* **33** (4) (2005) 644–663. School of Economics, University of Nottingham, University Park, Nottingham NG7 2RD, UK; School of Sociology and Social Policy, University of Nottingham, NG7 2RD, UK; School of Economics, Peking University, Beijing, China 100871.

JEL classification: J31; J42; O15; P23

Keywords: China; Labor market; Unemployment; Panel data; Wage premium; Fixed effects estimations

* Corresponding author.
E-mail address: simon.appleton@nottingham.ac.uk (S. Appleton).

0147-5967/$ – see front matter
doi:10.1016/j.jce.2005.08.005

1. Introduction

During the onset of China's urban reforms in 1983, Deng Xiaoping signalled the pragmatic, *ad hoc*, and gradual nature of his approach by using the metaphor that he was "crossing a river by feeling for the stones" (*Mozhe shitou guo he*). One of China's greatest challenges was transforming its planned labor allocation system into a well-functioning labor market. After more than two decades have elapsed, we can assess the progress made in this transformation and whether urban China is evolving towards a labor market similar to that found in OECD countries. One way to evaluate such progress is to estimate the change in the returns to productive worker characteristics and to unproductive worker characteristics during the transition. The evidence from the late 1980s and 1990s has been mixed. On the one hand, a rise in returns to education is often taken as a sign of a more competitive market that rewards workers for their productive skills. However, Knight and Song (2003) and Appleton et al. (2004) find apparently discordant trends in the rise of both gender wage differentials and the premium for Communist Party membership. In this paper, we investigate the impact of worker characteristics on wages in a systematic way that is consistent across time using a recent survey taken in 2002.

In addition, we examine the impact of ownership restructuring on the labor market, in particular, on the wage structure. Much of the impetuous for marketization came initially from the desire to provide greater autonomy to state owned enterprises (SOEs) resulting in a radical program of retrenchment within SOEs. Furthermore, liberalization of restrictions on rural–urban migration raise the possibility of workers with urban registration (*hukous*) having to compete for work with lower cost labor from rural areas. Generally, a more welcoming official attitude toward private enterprise generated growth of employment in private firms and individual businesses. Increased foreign direct investment and joint ventures also fostered a more competitive industrial structure. At the same time, the beleaguered urban collective sector has shrunken dramatically under pressure from both larger-scale SOEs and leaner private enterprises. Hence, China appears to be evolving towards a standard mixed economy with a sizeable state sector and a vibrant private sector.

Equally important to these changes in ownership is the gradual evolution of the industrial structure in China resulting in a relative decline in the importance of primary production and heavy industry. Despite the increase in manufacturing exports from China, an overall shift away from manufacturing into services has occurred. These developments are likely to undermine the position of the industrial worker and elevate the situation of white collar workers in tertiary sectors. However, apparently idiosyncratic features remain in the urban Chinese labor market. The coincidence of rising mass unemployment and rapid increases in real wages in the late 1990s is contrary to the predictions of competitive labor markets and to the empirical regularity of the wage curve observed in many countries. Similarly, the wage premium enjoyed by Communist Party members has not withered away, which suggests the continued influence of political forces of loyalty, power, and patronage on the rewards for labor. However, recent evidence raises doubts that these seemingly aberrant features of the Chinese labor market will endure.

The structure of the paper is as follows. Section 2 provides a brief account of the data and the econometric methods to highlight the comparison of cross-sectional and panel estimates of the wage structure in China. Section 3 reports the estimated results and draws implications about the change in returns to productive and unproductive characteristics of workers. In Section 4, we investigate the impact of the restructuring of Chinese industry on wages, including changes in industrial and occupational structure as well as ownership structure. Section 5 uses panel data estimates to provide evidence on the existence of a wage curve in China and the demise of the

Communist Party wage premium. Section 6 concludes by a summarizing the results and assessing the extent to which China has crossed the river.

2. Data and methods

We use three alternative econometric approaches to analyze changes in the pattern of wages in urban China. The first is a conventional estimation of wage functions using repeated cross sections. We estimate semi-logarithmic functions for wages of the form given in Eq. (1) separately for each year t. The dependent variable is the log wage per eight-hour period of work, $\ln W$, for each individual worker i. We specify:

$$\ln W_{it} = \beta_t' X_{it} + \varepsilon_{it} \tag{1}$$

where X is a vector of explanatory variables, β_t the corresponding coefficients, which may vary by year t, and ε is a random error term.

The cross sections are taken for the years 1988, 1995, 1999 and 2002 based on urban household surveys conducted as part of the China Household Income Project. The surveys were designed by a team of international scholars including the authors and researchers at the Institute of Economics and at the Chinese Academy of Social Sciences. Sub-samples were drawn from the larger annual national household income survey of the National Bureau of Statistics (NBS). The sub-samples cover 10 out of 31 provinces in 1988, 11 in 1995, 6 in 1999 and 12 in 2002. The questionnaires designed for the Household Income Project are more detailed than those in the official income surveys, particularly with respect to the measurement of income and labor issues. For the cross-sectional analysis, we construct a real wage variable that includes bonuses, price subsidies, which were important in 1988 before being largely withdrawn, regional allowances for working in Tibet or in mountainous areas, income in-kind, and income from secondary jobs.[1] Results from the 1988 survey are in Griffin and Zhao (1993) and those from the 1995 survey are in Riskin et al. (2001).

These surveys cover only households with urban registration (*hukou*). Consequently, we exclude rural–urban migrant households because they are denied urban *hukou* status. However, estimating wage functions of urban residents separately from those of migrants is appropriate because administrative controls make it extremely difficult for people of rural origin to acquire an urban *hukou* so that any sample selection bias is likely to be negligible. Confining the analysis to the sub-population having the urban *hukou* allows us to examine the changes in the wage structure for a specific group of people so that we may draw inferences about corresponding changes in economic well-being. Nonetheless, we are omitting an important dimension of the urban labor market by not being able to include migrants. Moreover, the importance size of this omission has increased over time with the sharp increase in rural–urban migration during the period.[2] Controls over rural–urban migration were loosened significantly in 1988 when the government allowed farmers to conduct business in cities, as Linge and Forbes (1990) discuss. The rise in

[1] Although fairly comprehensive, our wage variable does exclude some non-monetary benefits, e.g., pension accruals, health insurance, and housing. The contributions of these variables may vary under differing forms of ownership and over time. Nominal wages are converted into real wages by deflating by regional urban CPIs.

[2] Good data on the extent of rural–urban migration in China are not available because of the same constraints. However, official estimates of the floating-population, i.e., those living outside of their registration area, report an increase from around 2 million in 1983 to 61 million in 2000, as Fleischer and Yang (2004) discuss. Although not all of these people will be workers in urban areas, the majority will and the official estimates are considered to be conservative.

rural–urban migration is likely to have affected particular groups of urban workers differentially. Specifically, rural–urban migration is likely to have had a moderating impact on the wages of urban residents having similar characteristics as, or working in similar sectors to, migrants. Hence, the effect is greater on urban workers with less education and those working in the service and commercial sectors.[3]

Standard cross-sectional estimates may be biased by the failure to control for time-invariant unobserved determinants. Fixed effects estimates derived from panel data overcome this limitation.[4] All but the first survey included recall questions on individual wages in the four or five years preceding each survey. We use these recall data to form short panels on wages covering three episodes, namely 1990 to 1995, 1995 to 1999, and 1998–2002.[5] Asking workers to remember wages from up to four years ago is likely to lead to recall errors.[6] However, recall errors are limited by the fact that our households are sub-samples from larger surveys based on rotating three year panels surveyed by the NBS. The NBS requires respondents to keep written records on their income and consumption, which are scrutinized every ten days. As part of this exercise, an unidentified one third of our sample will have participated in such record keeping for the past three years. Respondents were also asked questions about job changes; hence, the reliability of recalled data on wages is improved.[7]

Nonetheless, fixed-effects estimators cannot be used to discern the impact of observed determinants of wages that are time invariant. This is a serious limitation for conventional wage functions because most of the personal characteristics that determine wages, e.g., gender and education, are effectively time invariant so that they cannot be included in a fixed-effects model of wages. Even experience, which does vary over time, provides no extra information if year dummies are included in the model. However, panel data on wages can be still be used to determine whether changes in the effects of these variables occur over time.

We consider Eq. (1) augmented with time-invariant individual-specific fixed effects, α_i, as follows:

$$\ln W_{it} = \beta_t' X_i + \alpha_i + U_{it} \tag{2}$$

[3] In the 1999 survey, the more-settled migrants were surveyed so that their characteristics are compared with those of workers having urban *hukou* in Table 1 of Appleton et al. (2004). Over half of the migrants were self-employed so that they do not compete directly for jobs with urban residents, only around 1% of whom were self-employed. Migrants tended to be less educated, younger and more likely to be male. The distribution of migrants across jobs was very different from that of urban residents, with a large concentration of migrant employment in service or retail sectors and relatively few migrants working as highly skilled or industrial workers.

[4] Although the fixed effects estimation eliminates bias from time-invariant unobserved factors having a proportional effect on wages, bias from factors influencing wages in other ways, e.g., additively or via slope-heterogeneity, are not eliminated.

[5] With the recall data on wages, we take only observations for which the worker was employed for the entire year. Years in which a worker is sacked or re-employed are dropped from the panel for data reasons. Since the survey does not provide recall data on months worked, we cannot infer a reliable monthly wage rate. Thus, we have an unbalanced panel in which not all workers in the sample are included for all years. We also drop observations for which the implied wages per day are less than eight *yuan* because such cases are unlikely to represent continuous full-time employment.

[6] If regarded to be classical measurement error in the dependent variable, any recall errors will reduce the statistical significance of the coefficients and, when using fixed effects estimation, bias them towards zero. Hence, the fact that we obtain significant results provides some assurance that recall errors are not too great.

[7] As a cross check on the validity of the recall data, we estimated a cross-sectional wage function for the 2002 sample using the log of recalled wages for 1999 as the dependent variable. The estimated coefficients on education, male, and Communist Party membership were with one or two percentage points of the corresponding results from the cross-sectional wage function estimated from the 1999 sample.

where U is a random error term. We estimate (2) using a fixed effects estimator by re-writing it in terms of interactions between the explanatory variables and the year dummies, D_{it}. For example, if the panel covers five years, we have:

$$\ln W_{it} = \gamma_1 X_i + \sum_{t=2}^{5} \gamma_t D_{it} X_i + \alpha_i + U_{it}. \tag{3}$$

The within-groups or fixed-effects estimator is based on:

$$\ln W_{it} - \overline{\ln W_{it}} = \sum_{t=2}^{5} \gamma_t \big[D_{it} X_i - \overline{D_{it} X_i}\big] + \big[U_{it} - \overline{U_{it}}\big]. \tag{4}$$

The fixed effects estimator removes potential biases caused by correlations between the explanatory variables, X_i, and the time-invariant unobserved characteristics of the individual, α_i. However, the cost is losing an estimate of the overall effect of the explanatory variables ($\beta_t = \gamma_1 + \gamma_t$) at any one point in time. This cost is less serious if we are interested primarily in the changes in the coefficients over time ($\beta_t - \beta_1 = \gamma_t$) rather than in the overall effect. The cross-sectional wage functions may imply different changes in returns over time from the panel ones for three reasons.

First, in comparing changes over time using the cross sections, we are using separate samples so that differences in the samples may yield different regression results. In contrast, with any one panel, a single sample is used. Second, the composition of the population may change over time, leading to variations in the relation between the observed determinants of wages, X, and the unobserved time invariant fixed effects, α. We refer to these changes as compositional or selectivity effects. For example, if an overall increase in schooling led to people with lower pre-existing ability receiving more education, the cross-sectional estimate of the returns to schooling might fall without contaminating the fixed-effects estimates. Similarly, if younger cohorts had higher quality schooling, the cross-sectional estimates of the returns to schooling would be affected. However, as a time invariant unobservable, school quality would be subsumed as part of the individual fixed effects and not affect the panel estimates of the change in the returns to schooling. Third, the panel estimates are likely to suffer more from measurement error, both because of the greater likelihood of recall error and because the within-groups estimator will amplify the effect of any measurement. Measurement error in the dependent variable reduces the likelihood of finding statistically significant results.

For data reasons, the third modelling exercise is confined to only the 1998–2002 panel based on recall data from the 2002 survey. We focus on estimating the impact of time-varying observed determinants of wages. Hence, we estimate equations of the following form:

$$\ln W_{it} = \beta' X_{it} + \alpha_i + U_{it} \tag{5}$$

where U is a random error term. Equation (5) differs from Eq. (2) in that the explanatory variables are time varying and because we estimate a single time invariant set of coefficients rather than allow the coefficients to change over time.

Using the 2002 panel, we construct three time-varying determinants of wages. First, the 2002 questionnaire inquired about whether the worker's enterprise had undergone a change in ownership due to restructuring. Hence, we construct time-varying dummies for the ownership sector in which an individual worked and derive a fixed-effects estimate of the wage premium, if any, for working in the state sector. Unlike the cross-sectional estimates, this fixed-effects estimate should be free from any biases due to the selectivity of employment in the state sector. However, it may be biased downwards due to an amplified measurement error in the independent variable.

The second time-varying determinant of wages is a proxy for the provincial unemployment rate. In many other countries, a significant negative effect of unemployment is found on wages, referred to as wage curve by Blanchflower and Oswald (1995). Unfortunately, we are not able to create a panel data set on province-level unemployment rates due to difficulties in accessing official administrative data and to the absence of recall questions on employment status in our surveys.[8] However, we create a proxy for unemployment rates by using the survey information on income in the previous five years. If people in the labor force in 2002 were reported as earning an annual income that equates to less than eight *yuan* per day, we considered these individuals to be unemployed.[9] This proxy is imperfect because it may include as unemployed some part-time workers; however, in China, such workers tend to be people who are unable to find work for the full year rather than workers who choose to work reduced hours or days each week in the year. Our proxy also measures unemployment relative to the labor force having urban *hukou* rather than including migrants, virtually none of whom are unemployed according to a sample of the more-settled migrants from the 1999 survey.

The third time-varying determinant from the 2002 data utilizes a question about when people joined the Communist Party, or some other political party. We take party membership as a time-varying determinant of wages in the 1998–2002 panel. By using a fixed-effects estimator, we can separate the causal effects of Communist Party membership from any selectivity effects of party membership. For example, the Communist Party may select particular able or hard-working people to become members, which would give the appearance of a wage premium for party membership in the cross sections regardless of whether party membership did tend to lead to higher wages. However, such selectivity would not contaminate the fixed-effects panel data estimates of the structure of wage changes irrespective of joining the Party or not. Despite its advantages, this estimation may still be subject to reverse causality, e.g., workers picked for promotion may be required to join the Party so that wage increases are more a cause than a consequence of Party membership, as Walder (1995) discusses.

3. Changes in returns to personal characteristics

We estimate the change in the returns to productive worker characteristics and unproductive characteristics using both the first and second approaches, namely cross-sectional wage functions and panel estimates of changes in the wage structure over time. Table 1 reports the results of the cross-sectional Mincerian wage functions estimated for 1988, 1995, 1999 and 2002. Following Mincer (1974), we regresses the log of wages on years of schooling, experience, and experienced squared. Only a few other control variables are included, specifically, dummies for gender, Communist Party membership, non-Han ethnicity, and province. Table 2 provides the results from a fuller specification in which we also control for job characteristics in terms of ownership sector, occupation, and industrial sector. Table 3 reports the results for the three panels on recalled

[8] Even if we had access to official data on unemployment rates by province, they may be unreliable. Wu (2004) blames lack of reliability for his failure to find a wage curve using official data on adult unemployment, although he does find one using data on youth unemployment.

[9] Less than 1% of the sample of those reported to be in full-time employment in the 2002 survey reported earning wages of 8 *yuan* per day. We suspect that these cases reflect reporting errors, i.e., they did not actually work full-time for the whole year, rather than genuinely low wages and excluded them as outliers from the earlier wage functions.

Table 1
Cross-sectional estimates of Mincerian wage functions

	1988	1995	1999	2002
Male	0.115	0.140	0.198	0.173
	(17.98)***	(11.21)***	(11.57)***	(15.76)***
Experience	0.047	0.060	0.038	0.026
	(40.29)***	(22.85)***	(11.39)***	(13.25)***
Experience squared term	−0.001	−0.001	−0.001	−3.35E-04
	(−25.48)***	(−16.6)***	(−8.64)***	(−34.1)***
Full-time education in years	0.036	0.056	0.067	0.075
	(29.59)***	(25.04)***	(19.71)***	(34.13)***
CP member	0.068	0.146	0.181	0.152
	(9.44)***	(11.05)***	(9.68)***	(12.07)***
Non–Han ethnicity	−0.006	−0.112	0.004	−0.006
	(−0.39)	(−3.72)***	(0.09)	(−0.23)
Constant term	1.763	1.970	2.295	2.724
	(72.09)***	(43.43)***	(36.48)***	(62.73)
No. of observations	17,733	12,245	6281	9791
Adjusted R-squared	0.34	0.24	0.24	0.294

Notes. (i) The dependent variable is log wage rate. (ii) T-ratios are in brackets. (iii) Provincial dummy variables are included in the models but their coefficients are not reported.
*** Statistical significance at the 1% level.

wages covering 1990–1995, 1995–1999, and 1998–2002. For these panel data models, we estimate Mincerian equations but allow the returns to personal characteristics to vary from year to year. Appendix Table 1 contains the means of the variables used in the cross sections.

In our data, two worker characteristics are prima facie productive, namely, education and experience. The Mincerian returns to education have been rising steadily in urban China.[10] From the cross-sectional evidence, the coefficient on years of schooling is 3.6% in 1988, which is low compared to other countries.[11] However, returns to schooling increase from 5.6% in 1995 to 6.7% in 1999 and to 7.5% in 2002. The panel evidence is broadly consistent with these cross-sectional results and suggests slightly larger increases in the returns to schooling over time. The fixed-effects panel estimates show an increase in returns to schooling of 2.7% between 1990 and 1995, followed by a further rise of 1.3% between 1995 and 1999, and an increase of 1.8% between 1999 and 2002. In addition, the panel data show that these increases have been smooth, rising year-on-year throughout the period from 1990 to 2002.[12] A comparison of the estimated returns to education in the Mincerian models in Table 1 with the fuller models in Table 2 indicates that roughly one third of the effect of education is attributable to a few job characteristics, namely,

[10] The Mincerian return to education is the coefficient on the years of education in a semi-log wage function. This interpretation of the coefficient depends on various restrictive assumptions, i.e., no pecuniary costs to education, the wage is the opportunity cost of labor, and workers are infinitely lived. Technically, this coefficient is the pure wage differential associated with a year of education.

[11] Psacharopoulos (1994) surveys the global literature and reports an average figure of 10%. Trostel et al. (2002) provide an arguably more reliable estimate of 4.8% for men and 5.6% for women from a twenty-eight country study. The Trostel et al. estimates are based on first-hand analysis of comparable data and do not suffer from a possible upwards publications bias in the literature survey.

[12] Education is entered interacted with year dummies in the panel data so that the associated t-ratios provide a test of whether the increase in the returns to education over time is statistically significant, which they are.

Table 2
Cross-sectional estimates of full wage functions

	1988	1995	1999	2002
Male	0.096	0.131	0.186	0.149
	(15.33)***	(10.93)***	(11.42)***	(14.34)***
Experience	0.044	0.058	0.043	0.027
	(39.53)***	(23.43)***	(13.10)***	(14.50)***
Experience squared term	−0.001	−0.001	−0.001	−3.54E-04
	(−24.58)***	(−17.76)***	(−10.72)***	(−9.07)***
Full-time education in years	0.028	0.033	0.039	0.049
	(20.62)***	(14.18)***	(10.82)***	(21.43) ***
CP member	0.052	0.075	0.098	0.089
	(6.89)***	(5.60)***	(5.25)***	(7.25) ***
Non–Han ethnicity	0.001	−0.103	0.017	−0.012
	(0.05)	(−3.54)***	(0.42)	(−0.48)
Ownership (*default variable is S.O.E.*)				
Urban collective	−0.142	−0.253	−0.180	−0.227
	(−18.45)***	(−14.41)***	(−6.60)***	(−10.67)***
Private enterprises	−0.341	−0.494	−0.038	−0.096
	(−2.86)***	(−7.17)***	(−0.89)	(−4.44) ***
Foreign-owned or joint venture	0.058	0.133	0.323	0.234
	(0.63)	(2.49)**	(5.53)***	(6.36) ***
Other ownership	−0.517	−0.299	−0.300	−0.042
	(−5.48)***	(−6.32)***	(−4.38)***	(−2.30)***
Occupations (*default variable is white-collar*)				
Private enterprise owner	0.040	0.109	0.147	0.067
	(0.67)	(1.57)	(1.78)*	(2.05)**
Blue collar	−0.061	−0.157	−0.144	−0.137
	(−7.89)***	(−10.64)***	(−7.30)***	(−10.21) ***
Other occupations	−0.175	−0.215	−0.271	−0.472
	(−1.59)	(−8.44)***	(−3.92)***	(−23.95) ***
Industry (*default variable is manufacture*)				
Primary industries	0.060	0.049	0.105	0.037
	(4.28)***	(1.25)	(2.73)***	(1.19)
Construction	0.018	0.005	0.100	0.040
	(1.19)	(0.16)	(2.63)***	(1.31)
Transportation and comm.	0.020	0.062	0.312	0.154
	(1.52)	(2.05)**	(10.82)***	(7.41) ***
Commerce	−0.010	−0.056	0.090	−0.097
	(−1.11)	(−3.00)***	(2.85)***	(−6.12) ***
Real estate	−0.078	−0.048	0.221	0.148
	(−3.30)***	(−1.57)	(7.34)***	(3.03) ***
Social welfare	−0.027	0.087	0.340	0.197
	(−2.26)**	(3.38)***	(10.36)***	(8.03) ***
Education	−0.061	0.134	0.306	0.215
	(−5.40)***	(7.01)***	(11.01)***	(10.73) ***
Sciences and research	−0.024	0.167	0.309	0.293
	(−1.52)	(5.05)***	(7.23)***	(7.38) ***
Financial sectors	−0.036	0.289	0.435	0.210
	(−1.61)	(7.34)***	(8.91)***	(6.43) ***
Government	−0.097	0.050	0.276	0.122
	(−8.65)***	(2.90)***	(10.09)***	(6.52) ***
Other industries	−0.153	−0.291	0.023	0.038
	(−3.21)***	(−6.59)***	(0.43)	(1.00)
Constant term	1.876	2.323	2.540	3.107
	(69.24)***	(48.04)***	(36.03)***	(66.65) ***
No. of observations	17733	12245	6281	9791
Adjusted R-squared	0.38	0.31	0.31	0.39

Notes. Same as for Table 1.

* Statistical significance at the 10% level.

** Idem., 5%.

*** Idem., 1%.

Table 3
Fixed-effects estimates of changes in returns to personal characteristics

	1990–1995	1995–1999	1998–2002
Male $(t+1)$	0.002	0.014	0.002
Male $(t+2)$	0.004	0.028**	−0.011
Male $(t+3)$	0.009	0.032***	−0.008
Male $(t+4)$	0.003	0.034***	−0.030**
Male $(t+5)$	0.010		
Experience $(t+1)$	0.000	−0.002	0.001
Experience $(t+2)$	0.001	−0.009***	−0.005**
Experience $(t+3)$	−0.001	−0.017***	−0.010***
Experience $(t+4)$	−0.008***	−0.017***	−0.015***
Experience $(t+5)$	0.009***		
Experience^2 $(t+1)$	−9.00E-06	3.69E-05	−3.78E-05
Experience^2 $(t+2)$	−4.34E-05	1.64E-04***	8.23E-5**
Experience^2 $(t+3)$	−2.02E-05	2.96E-04***	1.75E-04***
Experience^2 $(t+4)$	1.09E-04***	2.71E-04***	2.68E-04***
Experience^2 $(t+5)$	−2.79E-04***		
Education $(t+1)$	0.001	0.002	0.005***
Education $(t+2)$	0.001	0.005**	0.009***
Education $(t+3)$	0.007***	0.006**	0.014***
Education $(t+4)$	0.011***	0.013***	0.018***
Education $(t+5)$	0.027***		
CP member $(t+1)$	0.003	0.018	0.009
CP member $(t+2)$	0.007	0.067***	0.030***
CP member $(t+3)$	0.025**	0.081***	0.034***
CP member $(t+4)$	0.048***	0.073***	0.027***
CP member $(t+5)$	0.080***		
Non–Han $(t+1)$	−0.006	0.032	0.014
Non–Han $(t+2)$	0.016	0.044	0.014
Non–Han $(t+3)$	0.051***	0.051*	0.044**
Non–Han $(t+4)$	0.098**	0.087***	0.039*
Non–Han $(t+5)$	0.027		
Year $(t+1)$	0.068***	−0.043	0.049
Year $(t+2)$	0.142***	0.022	0.139***
Year $(t+3)$	0.122***	0.183***	0.220***
Year $(t+4)$	0.130***	0.261***	0.396***
Year $(t+5)$	−0.150***		
No. of observations	68,204	31,852	49,078
Within groups R-squared	0.061	0.095	0.226

Notes. (i) The dependent variable is log-wage rate. (ii) The independent variables are all interacted with year dummies, e.g., Male $(t+1)$ denotes a dummy variable for male interacted with the year dummy for $t+1$. (iii) The notation t refers to the base year of each panel, i.e., 1990, 1995 and 1998.

* Statistical significance at the 10% level.
** Idem., 5%.
*** Idem., 1%.

occupation, industrial sector, and ownership sector. Nonetheless, Table 2 shows a sustained rise in the returns to education even after controlling for these factors.

The increase in the returns to education is consistent with rewarding productive worker characteristics more and, thus, with the hypothesis of an increasingly competitive labor market. We find some support for this interpretation in the low return to schooling in China in 1988 compared with countries having more competitive labor markets. However, the rise may reflect other factors; in particular, skill-biased technological change may have increased the demand for skilled labor and thus the returns to education, as Krueger (1993) discusses for the US. The rival explanation for the increase in returns to education in the US is globalization and increased trade but this is less immediately applicable to China. Heckscher–Ohlin theory predicts that, although globalization should increase the returns to skills in developed countries, it should lower them in developing countries according to Wood (1994). However, for China, increased trade may complement both the institutional and the technological explanations. Greater openness puts more pressure on firms to set wages according to productivity and is also likely to lead to greater importation of skill-biased Western technology and methods.

The increase in the returns to schooling in China is larger than the increases in OECD countries during the same period. Vernon (2002) found only a moderate rise in the returns to schooling in the US, increasing from 9.5% in 1992 to 10.3% in 2000. Several other OECD countries, including Germany and France, experienced no increase in returns to education. Trostel et al. (2002) find no increase over time on average for returns to schooling in a pooled data set for twenty-eight countries. By contrast, Fleisher et al. (2005) find that most transition countries experienced marked increases in the returns to education, with the possible exception of Russia, during this period.

In addition to demand-side explanations, supply-side considerations also affect the returns to schooling in China. The liberalization of controls on migration increased the supply of predominantly low-skilled workers in urban areas, which may have moderated the increase in their wages and, thus, raised the returns to education. However, the degree of job competition that urban residents face from rural–urban migrants can be overstated. According to the 1999 migrant survey, most rural–urban migrants were self-employed and those in employment were distributed across industrial sectors differently from urban residents. Since the increase in the returns to schooling is apparent in the panel data, it is not likely to be attributable fully to an increase in the quality of education, which would affect cross-sectional estimates of the return to schooling. As the average educational quality of the workforce changes over time, a compositional effect occurs. However, the returns to education for a given individual with a given quality of education would not change. Hence, an increase in the quality of education cannot explain the rise in the returns to education in the fixed-effects estimates from the panel data.[13]

In addition to education, experience is the other worker characteristic predicted to raise productivity. Evaluating the change in the returns to experience is complicated by the inverse U-shaped relation between experience and wages. Figure 1 plots these relations for the cross section wage functions from Table 2. Up to 35 years of experience, the dominance relations between the wage-experience curves for the different cross sections are unambiguous. Within this range, experience was most highly rewarded in 1995, followed by 1988, then 1999, and finally 2002. However, after 35 years, the curves cross. The curves peak later in the estimates for 1988 and 2002 than in those for the 1990s. Consequently, workers with a lot of experience were rewarded more in 1988 compared to 1995 and also in 2002 compared to 1999. Despite these complications, the returns to experience tend to increase between 1988 and 1995 and to fall thereafter eventually to below their 1988 value. For example, the wage premium that a worker with 20 years of experience, which is roughly the average for the samples, was predicted to enjoy over a

[13] Since the increases in the returns to education are slightly higher in the panel than in the cross section, educational expansion during the period may have lowered somewhat the quality of the education possessed by entrants to the labor market.

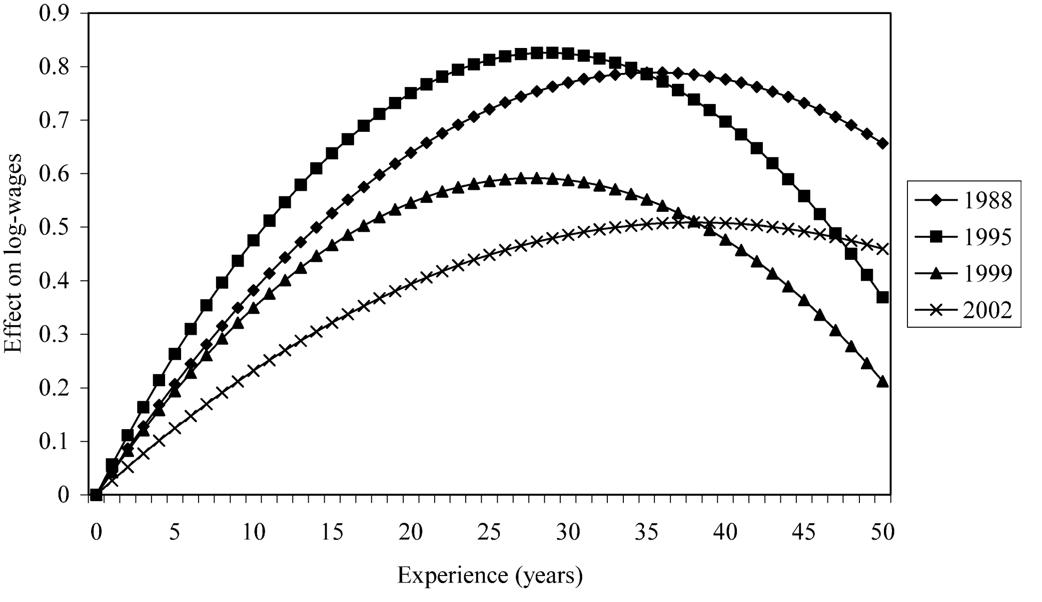

Fig. 1.

worker with no experience, ceteris paribus, stood at 89% in 1988, rose to 112% in 1995, and fell to 48% in 2002. The panel data estimates are consistent with these cross-sectional results. The dramatic decrease in the returns to experience is in contrast to the increase in returns to the other productive characteristic, education.

One possible explanation for the decrease in the returns to experience is that, unlike education, experience was over-rewarded prior to the reform in that payments for seniority were a central feature of the pre-reform wage structure. Regarding comparisons with countries having well-functioning labor markets, Meng and Kidd (1997) report estimates of returns to experience in the UK in 1980, in the US in 1989 and in Australia in 1990 that are lower than our estimates for China in 1988, 1995 and 1999. Only in 2002, do the returns to experience in urban China become approximately equal to those in the UK and Australia and less than those in the US in the years specified for these countries. Moreover, Appleton et al. (2002) find an inverse-U relationship arose between general work experience and the probability of retrenchment in China in 1999. If experience was over-rewarded in the pre-reform period, experienced workers would be most at risk from retrenchment by enterprises attempting to increase profitability so that their wage premiums would decline after 1995. In general, a higher level of job insecurity may have exerted downward pressure on the wages of experienced workers.

Regarding unproductive worker characteristics, increased competition is expected to reduce the returns attributed to them during the pro-market reforms. We identify three worker characteristics that are prima facie unproductive, namely, gender, political affiliation, and ethnicity. Gender differences in wages are often attributed to discrimination so that the increase in the gender wage gap in the 1990s appears to contradict the notion of an increasingly competitive labor market. However, gender differences in wages, even after controlling for education, experience, and other observed characteristics, are a pervasive feature of labor markets in OECD countries. The cross-sectional Mincerian wage functions for China in 1988 imply that men earned 12%

more than women, ceteris paribus.[14] This pure gender gap rose to 15% in 1995 and to 22% in 1999. International comparisons suggest that the administered pay schedules used in the planning period may have been more gender equitable than those in market systems. For example, Vernon (2002) estimates that the pure gender pay gap in the US rose from 21% in 1992 to 25% in 2000. The results for China indicate that the increase in the gender wage gap was halted and even reversed as the estimate for 2002 is 19%.

In addition, the panel data show significant increases in the pure gender gap only for the period from 1995 to 1999. The wage premia for men are not statistically significant in the 1990-to-1995 panel and the last year of the 1999-to-2002 panel indicates a significant drop in this male premium. Compositional effects may account for some of the increase in the wage premium for men between 1990 and 1995 and from 1999 to 2002. An increase in women's share of employment from 40% in 1988 to 47.5% in 1995 may have lowered the average unobserved productivity of female employees. When women's labor force participation was lower in 1988, a sample selectivity effect by which women with higher wage potential dominated the workforce may have influenced the results. This explanation is consistent with the practice of discrimination against women, indicating that the extent of wage discrimination is understated when only more productive women are employed. By contrast, both the cross sections and the panel show a rise in the pure gender wage gap between 1995 and 1999 indicating that wage discrimination against women may have increased during that period. Appleton et al. (2002) show that women suffered more during retrenchment because they faced not only a higher risk of retrenchment but, if retrenched, they had a lower probability of re-employment than men.

Similar to the wage premium for male workers, the wage premium for Communist Party membership is a form of discrimination or favoritism because it is not justified by underlying productivity differentials. As did the male wage premium, the premium for Communist Party (CP) membership rises during the 1990s but no divergence between the cross-sectional and panel estimates occurs in this case. In the Mincerian cross-sectional estimates, the coefficient on the dummy variable for CP membership increases from 0.068 in 1988 to 0.181 in 1999. Although the panel estimates do not extend as far back as 1988, they indicate that the premium rose by 0.080 between 1990 and 1995 and by an additional 0.073 by 1999. This overall increase in the CP premium in both cases implies that a change in the composition of the party membership, for example, admitting more productive members, cannot explain this phenomenon. However, CP members may have more productive characteristics, which are not controlled for in the wage functions, so that the increase in their wage premium may reflect a general increase in the returns to productive characteristics.

After 1999, the cross-sectional and panel estimates of the returns to party membership diverge. The premium for CP membership decreases in the cross section estimates with the coefficient in 2002 at only 0.152, but it continues to increase by 0.018 between 1999 and 2002 in the panel estimates. One possible explanation is a change in the composition of party membership, with new entrants having unobserved characteristics that tend to be less productivity-enhancing than the characteristics of more-established party members. If this change in the unobserved productivity characteristics of party members is large enough, the cross-sectional returns would decrease even if the true benefit to party membership had risen as the panel data estimates indicate.

[14] Following Halvorsen and Palmquist (1981), we calculate percentage effects as equal to $\exp(\beta) - 1$, where β is the relevant coefficient in the log-wage functions.

Contrary to the experiences of ethnic minorities in other transition economies, e.g., the findings of Giddings (2003) for Bulgaria, we find little or no evidence of increased wage discrimination against ethnic minorities during China's urban reforms. Only in the cross section for 1995 are ethnic minorities paid significantly lower wages, ceteris paribus. In some OECD countries, notably the US, Mincer-type wage equations identify persistent and significant ethnic differences in wages even after controlling for education, experience, and other characteristics (Vernon, 2002). Moreover, interactions between a dummy for minority ethnicity and year dummies are positive in all three panels. Hence, during this period, urban workers from minority ethnic groups enjoyed faster increases in wages than other workers, ceteris paribus. The diverging results from the panel data and the cross sections may reflect the faster wage growth of urban-to-urban migrants. The cities in our surveys are all in ethnically Han areas; hence, any non-Han Chinese workers in our sample are more likely than their Han counterpart to have moved into the cities recently.[15] As Chiswick (1978) shows in the context of US immigration, migrants tend to have faster wage growth than natives. Although they may initially lack information about job-opportunities or firm-specific capital, migrants catch up with local workers as their stay lengthens. In addition, migrants may be selected for their particularly high abilities.[16] Hence, a higher potential for earnings growth by non-Han migrant workers could explain the time trends observed in the panel data. However, these trends could be mitigated in the cross-sectional estimates by a steady flow of initially disadvantaged new non-Han migrants into urban areas.

4. The effect of enterprise restructuring on wages

During the period from 1988 to 2002, profound shifts in industrial and enterprise structure occurred in urban China as the size of the private sector increased while employment in foreign-owned or joint-venture enterprises expanded. As Appendix Table 1 reports, only 0.8% of those in our sample worked in private enterprises in 1988 but that percentage had risen to 11.3% by 2002.[17] The proportion of workers employed in SOEs, or in other public sector employment, fell during the period. However, the biggest contraction in employment occurred in urban collectives with a decrease from 20% in 1988 to 5.9% in 2002. Contemporaneously with ownership restructuring was a gradual shift in industrial structure. The share of workers in primary and secondary sectors fell from 50.3% in 1988 to 42.0% in 2002. Services, in particular, commerce and government departments, gained employment share. These changes are reflected in the occupational composition of the Chinese urban workforce. Whereas blue collar workers outnumbered white collar workers in 1988, the situation was reversed in the 1990s. By 2002, industrial workers accounted for slightly over one third of all employees in urban China. The changes in ownership and industrial structure were linked with retrenchment in unprofitable SOEs in heavy industry.

The pressures that led to changes in the composition of employment may have affected wages. Declining sectors are more likely to be loss-making and the increased autonomy given to enterprises in setting wages suggest a negative impact on the wages received by workers in those

---

15 Since our sample consists of workers with urban *hukou*, we consider only urban-to-urban migration. However, talented individuals from rural areas may have acquired the urban *hukou* by virtue of being selected for a university education.

16 Chiswick (1978) explains his observation of high earnings growth of US immigrants in the 1950s and 1960s by positive selectivity.

17 The rise in private sector employment is likely to be understated because the percentage of individuals whose ownership sector could not be determined increased from 0.9% in 1988 to 8.9% in 2002.

sectors.[18] The cross-sectional estimates in Table 2 show a widening gap between pay in the SOE sector and in the withering urban collectives, mainly between 1988 and 1995. By contrast, differentials between wages in SOEs and private or foreign enterprises change mainly after 1995. In 1988, workers employed in private enterprises are estimated to have been paid 29% less than their counterparts in SOEs. This gap widens between 1988 and 1995, but narrows during the period of retrenchment in the public sector. By 2002, this differential is only 9%. The wages of employees in foreign-owned or joint venture enterprises are not significantly different from those of workers in SOEs in 1988, but are estimated to be 26% higher by 2002.

Using information on ownership restructuring from the 2002 survey, we include the ownership sector as a time-varying determinant of wages in a fixed-effects model estimated using the 1998–2002 panel. Unlike the model in Table 3, this model does not allow for changes in the returns to personal characteristics. However, several other time-varying explanatory variables are included, in particular, an estimate of the local unemployment rate and dummies for membership in political parties. In principle, the resulting fixed-effects estimates of the influence of the ownership sector can be compared with those in cross section regressions for 1999 and 2002 in Table 2. As Table 4 indicates, the panel estimates of the impact of ownership sector tend to be generally smaller in absolute size than the cross-sectional estimates. For example, according to the panel estimates, working for an urban collective rather than an SOE does not affect signifi-

Table 4
Fixed-effect wage function with time-varying variables: 1998–2002 panel

Variable		Coefficient	T-ratio
Unemployment rate		−0.667	−6.01***
Communist Party membership		0.054	4.86***
Other political party membership		0.041	1.76*
Ownership (default variable is state owned)			
Urban collective ownership		−0.014	−0.77
Urban private ownership		−0.062	−2.32**
Urban individual business		−0.080	−1.95*
Joint venture		0.054	1.82*
Foreign direct investment		−0.059	−0.78
State-control share-holding		0.022	1.58
Other share-holding		−0.026	−2.57***
Government department/public agents		−0.043	−0.21
Other		−0.034	−1.60
Year dummy 1999		0.055	14.75***
Year dummy 2000		0.110	22.64***
Year dummy 2001		0.164	27.63***
Year dummy 2002		0.263	39.50***
Constant		3.546	181.55***
Number of observations	46,829		
R-sq. within	0.207		
Dependent variable	log hourly wage rate		

* Statistical significance at the 10% level.
** Idem., 5%.
*** Idem., 1%.

[18] Knight and Li (2002) provide evidence of a link between enterprise profitability and wages in China.

cantly a worker's wage. By contrast, the 2002 cross section estimates imply that wages would be lower by one fifth for urban collective workers.

Such discrepancies may be due to differences in either firm or worker unobserved characteristics. The cross-sectional estimates are likely to be affected by differences in firm size and capital-worker ratios, neither of which is available in the household surveys. However, despite restructuring, these variables may not change very rapidly over time so that they have less impact on the fixed-effects estimates. The discrepancy between the two sets of estimates may also reflect sorting by workers with unobserved productive characteristics into higher paying ownership sectors. Suppose that workers in urban collectives had pre-existing unobserved characteristics that resulted in them being paid less than workers in SOEs; in which case, the ownership sector would have no causal effect.

The cross-sectional and panel data estimates are consistent with respect to some effects of private ownership on wages. In all cases, we find a significant negative effect implying that, on average, the default sector, i.e. SOEs, overpays compared to the private sector. In the cross-sectional estimates, the private sector pays 3.9% less in 1999, ceteris paribus, and this differential increases to 9.2% in 2002. In the 1998–2002 panel, the private-sector differential is estimated to be 6%. This broad agreement between the cross-sectional and panel estimates suggests that selectivity is not a problem for comparing wages in the two sectors. The results suggest that retrenched workers have an incentive to seek re-employment in the SOE sector rather than in the private sector.

The panel estimates provided mixed results on the effect of working for joint ventures or companies having foreign direct investment. We find a significant wage premium for working in a joint venture, but no significant impact for workers in a firm having foreign direct investment.[19] However, the size of the premium for working in a joint venture is much lower than that observed in the cross sections, which may reflect only the ability of such enterprises to recruit higher caliber personnel. Moreover, these findings should be taken with some caution given the small size of the sectors and the resulting limited number of observations of movement into them within the 1998–2002 panel.

Regarding changes in industrial sector and occupation, the estimates from the cross-sectional models in Table 2 indicate that wages in many of the burgeoning service sector jobs rise relative to those in manufacturing, which is the default industrial sector in the regressions. A comparison of the results for 2002 and 1988 show an increase in the wage premium in all service sectors, except for commerce. Some service sectors, i.e., education, social work, and government administration, switch from paying significantly less than manufacturing to paying significantly higher wages. However, commerce workers were paid no differently from manufacturing workers in 1988 but, by 2002, they received 10% lower wages. The commerce sector has been affected most by the liberalization of rural–urban migration during this period. The influx of self-employed and employed migrants working in commerce moderated increases in the wages of urban residents employed in this sector.

In addition to these changes in industrial structure, occupational wage structure, broadly defined, has also changed during the period. We compare white-collar workers, i.e., those in professional and technical, managerial and administrative, or clerical posts, with blue-collar workers, i.e., those in industrial, commercial, or service posts. Between 1988 and 2002, our

[19] Citing Pearson (1992), a referee reports that pay in joint ventures was legislated to be 20% higher than in SOEs. However, Pearson covers the period up to 1988 only while our panel results suggest a substantially lower premium in the period from 1998 to 2002.

cross-sectional estimates in Table 2 indicate that the pure wage differential between these two groups more than doubled. Similarly, the wages of private-enterprise owners became significantly greater than those of white-collar workers during this period. In summary, although restructuring has brought the wage structure in urban China closer to its counterparts those in OECD countries, considerable unemployment has been generated. The impact of unemployment on wages is examined in the next section.

5. Idiosyncratic features of Chinese wage functions

We conclude our examination of the behavior of wages in urban China by presenting new evidence on two apparently idiosyncratic aspects, namely, the absence of a wage curve and the existence of a wage premium for Communist Party members.

Regarding the wage curve, retrenchment within the state sector was a key aspect of restructuring Chinese industry. The radical reform of *xia gang*, first experimented with in 1994 and finally launched fully in 1997, was intended to resolve the problem of inefficiency in the state sector by laying-off a quarter or more SOE workers within a four-year period from 1997 to 2000. As a consequence, urban unemployment rose sharply but, at the same time, real wages increased considerably raising doubts about the competitiveness of the Chinese labor market. Since the state sector was widely believed to have suffered from over-manning, wages should not have risen on account of this retrenchment policy. However, an increase in wages is consistent with a rent-sharing hypothesis in which retained workers capture a share of the wage bill that would otherwise have been saved by dismissing excess workers. Nonetheless, the phenomenon of rising wages and unemployment in the late 1990s in China is not consistent with the wage curve that presents a trade-off between real wages and unemployment found in other countries by Blanchflower and Oswald (1995).

Sabin (1999) uncovered evidence of a wage curve in China for an earlier period from 1980 to 1992, but this was before urban unemployment was a significant economic issue. Wu (2004) found no wage curve using official provincial unemployment data, but did uncover one using data on youth unemployment. We take a survey-based proxy for the provincial unemployment rate as a determinant of wages in the fixed-effects model estimated in Table 4. Using the conventional semi-log functional form, the coefficient for this variable is negative and statistically significant. Evaluated at the mean unemployment rate of 14.5%, the coefficient implies an elasticity of real wages with respect to unemployment of −0.046. This elasticity is low compared to values of around −0.1 found in OECD countries by Blanchflower and Oswald (1995). The lower elasticity in China may reflect the effect of measurement error in our unemployment variable. However, it does imply significant moderating pressure on wages in provinces where unemployment is higher. Hence, we find some evidence of a wage curve in China from 1998 to 2002.

Regarding Communist Party membership, our cross-sectional wage functions in Table 2 indicate a substantial wage premium, even after controlling for human capital and some characteristics of the enterprises. Previously, we interpreted this result as a payment for an unproductive characteristic and, as such, considered it to be evidence refuting the hypothesis of increased competition. However, workers with higher unobserved productivity may be more likely to be party members. The Communist Party has a tradition of targeting exemplary workers and, more recently, it may have attracted more ambitious workers wishing to use membership to advance their careers. Since the 2002 survey includes a question on when members joined the Party, we can construct a time-varying measure of party membership. Thus, using a fixed-effects estimator, we can purge our estimates of the impact of party membership of the influence of differences in time-invariant unobserved productivity.

The fixed-effects estimates reported in Table 4 indicate a 5.5% wage premium attributable to CP membership, when the latter is entered as a time-varying determinant in the wage function for the 1998–2002 panel. Although statistically significant, this estimate is substantially below the cross-sectional results for either 1999 or 2002. Hence, a large part of the apparent wage premium for CP membership reflects the return to unobserved time invariant factors. However, some caution is required in applying this result because the panel is rather short so that it may take more time for the benefits of party membership to appear.

The 2002 survey also provides information on membership in China's other political parties. These eight supposedly democratic parties have less than half a million members, compared to 67 million in the Communist Party in 2002, and traditionally have represented members who are supportive of the Communist regime but come from non-labor backgrounds, as Townsend (1967) discusses. When membership in these parties is entered as a time-varying variable in Table 4, the wage premium attributable to membership is similar to that for membership in the Communist Party. Perhaps membership in these parties makes one a fellow traveler and confers the same privileges and favoritism enjoyed by CP members. Alternatively, party membership may bring benefits through access to networks rather than through power relations per se. Membership in any political party in China, much like membership in the Rotary Club in other societies, may represent an investment in social capital.

6. Conclusion

With the emergence of a mixed economy and the growing dominance of services, China's urban labor market appears to be evolving towards a competitive market similar to those found in OECD countries. The rising share of workers employed in the private sector increases the extent to which pay is determined by productivity. The premium paid to state-sector employees is withering away. Based on changes in ownership due to restructuring between 1998 and 2002, we estimate the effects of ownership on wages, controlling for selectivity effects. Our results imply that some of the wage differentials by ownership reflect selectivity effects as more productive workers gravitate to enterprises with foreign investment while less productive workers stay in urban collectives. However, both panel and cross-sectional estimates yield a premium for being employed by the state rather than a private enterprise. Although this public sector wage premium is modest at around 5%, it is statistically significant.

The shift in employment from primary and secondary to tertiary service sectors in China is also mirrored in the urban wage structure. During the 1990s, this change in industrial structure ended the privileged status of blue-collar manufacturing workers. In addition, the gap between blue-collar and white-collar workers has widened. Some of China's recent industrial transformation results from the painful process of retrenchment in the state-owned sector, particularly in unprofitable heavy industry. Although rising unemployment in the late 1990s was accompanied by real wage increases, we find evidence of a relatively flat wage curve in China from the 1998–2002 panel as higher provincial unemployment exerts a moderating effect on wages.

To assess directly the extent to which China has moved towards a competitive labor market, we need information on worker productivity to see if it matches pay, evidence on barriers to job mobility, and data on whether employers are allowed to determine the optimal size and composition of their workforces. Lacking such information, we focus on wage data from four surveys of urban workers to investigate the changes in productive and unproductive characteristics over time. We find an increase in the returns to education in China that is consistent with an increasingly competitive labor market. However, international comparisons suggest that education was undervalued during the planning period so that we should expect it find such a result during the transition to a market-based system. Cross-sectional evidence from the recent 2002 survey confirms the trend observed in the earlier surveys of an increase in the returns to education. Short panels constructed from recall data on wages also show year-on-year increases in the returns to education from 1990 to 2002.

However, a different picture emerges from trends in the returns to experience and gender differentials. Although experience is commonly regarded as a productive characteristic, no a priori reason exists to regard it as being undervalued in the administratively determined wage system prior to the reform in China. However, payment according to seniority was a distinctive feature of wage-scales in China and international comparisons suggest that experience was overvalued prior to reform. Thus, our finding of a decrease in returns to experience in China is consistent with a move toward greater efficiency and a more competitive labor market. Moreover, by 2002, returns to human capital measured by both education and experience in China were closer to those observed in OECD countries.

Interpreting the increase in the gender wage differential in urban China is controversial. Higher wages for men compared to women with the same education and experience are commonly assumed to reflect discrimination rather than productivity. However, such differentials are persistent and almost universal features of economies with labor markets that are otherwise assumed to be well-functioning and competitive. International comparisons suggest that wages in urban China at the beginning of the transition were relatively gender equitable and, even though the pure gender differential has widened sharply since 1988, it remains smaller than that observed in the US. Similarly, any wage discrimination against ethnic minorities in the Chinese cities in our surveys appears prima facie to be less than its counterpart in the US. Moreover, we find no evidence that the transition has aggravated discrimination; rather, individual non-Han Chinese in our data report relatively rapid earnings growth.

Finally, we consider the paradox of the increasing premium to Communist Party membership in urban China during a period of pro-market reform and of a diminishing influence of the Communist Party on promotion. The recent 2002 data suggests that this benefit is withering away. Although the wage premium appears to be rising in our panel estimates, it has fallen in the cross-sectional estimates perhaps indicating a dilution of the unobserved productivity of new members or a vintage effect of party membership. Taking party membership as a time-varying determinant of wages, we find a much lower premium attributable to being a Communist. Thus, entrance to the party in the period 1998 to 2002 is associated with only a moderate rise in wages. These lower estimates compared to the cross-sectional results may reflect party membership signalling pre-existing higher productivity. Alternatively, the next generation of party members may be less able to secure the privileges enjoyed by those who entered at a time when the Communist Party and the state had a greater role in the allocation of labor. Moreover, membership in smaller supposedly democratic parties now confers benefits to membership in the Communist Party, perhaps indicating that the value of access to networks rather than political power relations per se.

Acknowledgments

We are grateful to the CCK Foundation for a research grant and the Ford Foundation for its support of the data collection. Useful comments on the papers were provided by Meng Xin, two anonymous referees, and participants at presentations at the Australian National University and the University of Gothenburg.

Appendix Table 1
Mean statistics of workers

	1988	1995	1999	2002
Log daily wage	2.76	3.08	3.25	3.72
Male	60.07	52.57	54.16	55.74
Experience	21.05	22.00	22.21	23.65
Full-time education in years	10.04	10.56	11.24	11.42
CP member	23.47	24.51	26.89	29.95
Non–Han ethnicity	3.77	4.31	4.17	4.21
Ownership sector				
State owned sector	77.68	79.03	77.46	72.05
Urban collective	20.28	15.06	13.20	5.93
Private enterprises	0.77	1.65	4.59	11.26
Foreign-owned or joint venture	0.37	1.28	1.99	1.90
Other ownership	0.92	2.98	2.77	8.86
Occupation				
Private enterprise owner	1.21	1.47	1.42	8.96
White collar	45.43	52.83	50.18	45.31
Blue collar	52.76	37.44	45.49	35.85
Other occupations	0.60	9.56	2.92	9.89
Industrial sector				
Primary industries	4.13	2.64	3.53	2.77
Manufacturing	42.73	39.86	31.73	36.23
Construction	3.41	2.87	4.34	2.972
Transportation and communication	6.75	4.86	9.25	7.22
Commerce	14.41	14.23	10.79	17.56
Real estate	2.45	3.80	10.64	1.07
Social welfare	4.55	4.38	4.42	5.13
Education	7.22	7.11	7.28	9.01
Sciences and research	2.89	2.27	2.18	1.69
Financial sectors	1.53	1.92	2.07	2.56
Government	8.43	11.31	8.82	11.98
Other industries	1.52	4.71	4.94	1.82

References

Appleton, Simon, Knight, John, Song, Lina, Xia, Qingjie, 2002. Labour retrenchment in China: Determinants and consequences. China Economic Review 13, 275–299.

Appleton, Simon, Knight, John, Song, Lina, Xia, Qingjie, 2004. Contrasting paradigms: Segmentation and competitiveness in the formation of the Chinese labour market. Journal of Chinese Economic and Business Studies 2, 185–206.

Blanchflower, Daniel, Oswald, Andrew, 1995. The Wage Curve. MIT Press, Cambridge, MA.

Chiswick, Barry R., 1978. The effect of Americanization on the earnings of foreign-born men. Journal of Political Economy 86, 897–921.

Fleischer, Belton M., Yang, D., 2004. China's Labor Market. Paper presented at the Conference on China's Market Reforms, Stanford Center for International Development, Stanford University, September 19–20, 2003.

Fleisher, Belton M., Sabrianova P., Klara, Xiaojun, Wang, 2005. Returns to skills and the speed of reforms: Evidence from Central and Eastern Europe, China and Russia. Journal of Comparative Economics 33, 351–370.

Giddings, Lisa, 2003. Continued decline for ethnic minorities in the transition?. Economics of Transition 11, 621–648.

Griffin, Keith, Zhao, Renwei (Eds.), 1993. The Distribution of Income in China. Macmillan & Co., London.

Halvorsen, Robert, Palmquist, Raymond, 1981. The interpretation of dummy variables in semilogarithmic equations. American Economic Review 70, 474–475.

Knight, John, Li, Shi, 2002. How does firm profitability affect wages in urban China? Mimeo. Department of Economics, University of Oxford, Oxford.

Knight, John, Song, Lina, 2003. Increasing urban wage inequality in China. Economics of Transition 11, 597–619.

Krueger, Alan, 1993. How computers have changed the wage structure: Evidence from microdata, 1984–1989. Quarterly Journal of Economics 108, 33–60.

Linge, Godfrey, Forbes, Dean, 1990. China's spatial development: Issues and prospects. In: Linge, Godfrey, Forbes, Dean K. (Eds.), China's Spatial Economy – Recent Development and Reforms. Panther Press, Hong Kong.

Meng, Xin, Kidd, Michael P., 1997. Labor market reform and the changing structure of wage determination in China's state sector during the 1980s. Journal of Comparative Economics 25, 403–421.

Mincer, Jacob, 1974. Schooling, Experience and Earnings. Univ. of Chicago Press, Chicago.

Pearson, Margaret, 1992. Joint Ventures in the People's Republic of China: The Control of Foreign Direct Investment Under Socialism. Princeton Univ. Press, Princeton.

Psacharopoulos, George, 1994. Returns to investment in education: A global update. World Development 22, 1325–1344.

Riskin, Carl, Zhao, Renwei, Li, Shi, 2001. China's Retreat from Equality: Income Distribution and Economic Transition. M.E. Sharpe, Armonk, New York.

Sabin, Lora, 1999. The development of urban labour markets—China's urban wage curve, 1980–92. In: Cook, Sarah, Maurer-Fazio, Margaret (Eds.), The Workers' State Meets the Market: Labour in China's Transition. Frank Cass, London.

Trostel, Philip, Walker, Ian, Woolley, Paul, 2002. Estimates of the returns to education in 28 countries. Labour Economics 9, 1–16.

Walder, Andrew G., 1995. Career mobility and the Communist Political Order. American Sociological Review 57, 524–539.

Wood, Adrian, 1994. North–South Trade, Employment and Inequality: Changing Fortunes in a Skill-Driven World. Clarendon Press, Oxford.

Wu, Zhongmin, 2004. Wage curve for urban China: A panel data approach. Applied Economics Letters 11, 425–428.