Economic Reforms and Gender Inequality in Urban China

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I. Introduction

The impact of economic reforms on the gender earnings gap in transition economies has been examined by several authors with mixed findings. For example, Brainerd (2000) finds that economic reforms increased the gender earnings gap in Russia and Ukraine but reduced it in most Eastern European countries, such as Hungary, Poland, Czech Republic, and Slovak Republic. Hunt (2002) shows that while the gender earnings gap fell in East Germany after German unification, almost half of it was attributable to the exit from employment of low-skilled women. Münich, Svejnar, and Terrell (2005) suggest that economic reforms narrowed the gap in the Czech Republic. Studies on China's gender earnings gap mostly find that economic reforms increased the gender earnings gap. For instance, Gustafsson and Li (2000) show that the gender earnings gap in urban China grew slightly, from 15.6% in 1988 to 17.5% in 1995. A similar pattern has also been documented by Knight and Song (2003) and Yang (2005). A recent study by Zhang et al. (2007) shows that the gender gap kept increasing even after 1995, from 14.7% in 1988 to 22.6% in 1995 and further to 27.2% in 2004. Nevertheless, compared with other transitional economies, the gender earnings gap in China is still relatively small.

Several factors might be responsible for China's small and relatively stable gender earnings gap. First, although China's economic reforms commenced in the late 1970s, early reforms mainly affected rural areas. The urban labor market was not affected until the mid-1990s (e.g., Meng 2004; Cai, Park, and Zhao 2008). This is evident from changes in the payroll employment share of state-owned units (SOU) and of collective-owned enterprises (COE). The payroll employment share of SOUs was 73.4% in 1988, 73.5% in 1995,

I would like to thank Jeffrey Smith, Yu Xie, and two anonymous referees for their helpful comments.
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and 63.1% in 2003.[1] This suggests that studies using pre-1995 data might not be able to reveal the impact of recent urban reforms. Second, as has been pointed out by Hunt (2002), the small increase in the gender earnings gap might be attributable to the decline in the employment rate of low-skilled workers. Because the average skill level of women is lower than that of men, a decline in the employment rate of low-skilled workers has a larger effect on women's than on men's employment, which narrows the gender skill gap. For example, Giles, Park, and Cai (2006) show that while the female labor force participation rate declined from 77.4% in 1996 to 66.3% in 2001, the male labor force participation rate fell by 6.7 percentage points from 93% to 86.3%. The decrease in labor force participation rate was more pronounced for low-skilled women. Zhang et al. (2007) show that the educational level of employed women increased by 2 years while the corresponding increase for men was 1.2 years. The evidence from both studies suggests that changes in the labor force composition of women could play a significant role in shaping the gender earnings gap in China during the reform period.

This study uses information extracted from the China Health and Nutrition Survey (CHNS) to tackle this issue. Compared with previous studies that largely used the 1988 and 1995 China Household Income Project (CHIP), the CHNS enables us to document the evolution of the gender earnings gap over a much longer period, between 1989 and 2004. By extending the sample coverage to after the 1995 period, we can compare the impact of recent reforms that mostly affected urban areas with the impact of earlier reforms. Our results are complementary to those of Zhang et al. (2007), whose sample mostly consists of households from relatively developed provinces. Combining the evidence documented in these two studies can depict a broader picture of the evolution of the gender earnings gap in China. In addition, since the CHNS is a longitudinal survey, we can use it to jointly examine changes in employment probability and changes in the gender earnings gap.

Consistent with previous studies, we find that the gender earnings gap in urban China rose from 13% in 1989 to 17% in 1997. However, unlike Zhang et al. (2007), we find that the gender earnings gap stabilized around that level for the rest of the sample period. This pattern is not sensitive to whether we focus on individuals between the ages of 16 and 60 or 16 and 45. It is also not sensitive to whether we use the entire sample or only focus on people who resided in households that were surveyed throughout the entire sample period.

[1] The information is extracted from table 1-14 of the China Labour Statistical Yearbook 2005. The payroll employment, which is called *zhi gong* in Chinese, only includes staff and workers. Hereafter, we will use "payroll employment" and "employment" interchangeably.

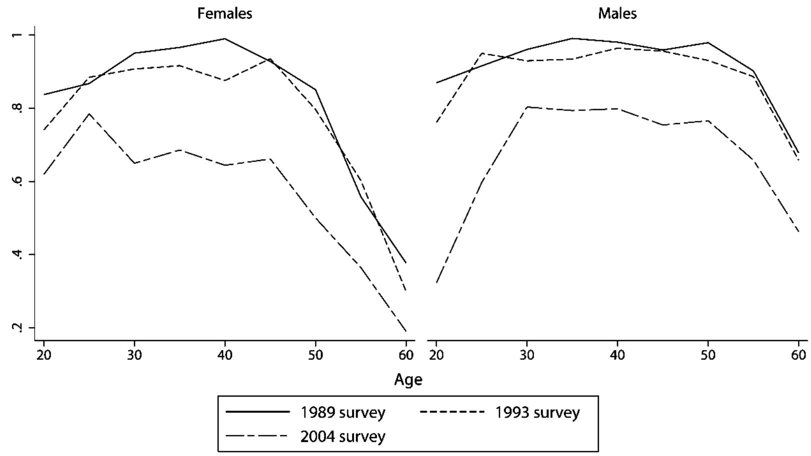

Figure 1. Employment rates in 1989, 1993, and 2004

Therefore, the documented pattern is unlikely to be driven by sample attrition or changes in the sample coverage of the CHNS data. The increase in the earnings gap in the earlier reform period was driven by a widening gap at both the top and bottom of the earnings distribution. Our quantile regression results show that the gap increased from 6.3% in 1989 to 10.5% in 1997 at the 10th percentile and from 13.2% in 1989 to 15.1% in 1997 at the 90th percentile. The seemingly stable gap between 1997 and 2004 was the result of a considerable widening gap at the bottom and a narrowing gap at the top of the earnings distribution. For example, the gap rose from 10.5% in 1997 to 17.7% in 2004 at the 10th percentile, but it fell from 15.1% to 6.9% over the same period at the 90th percentile. As a result, the mean gender earnings gap hardly changed. This evidence suggests that while the gender earnings gap in the early reform period was characterized by men gaining ground on women throughout the entire earnings distribution, high-wage women regained some ground in the later reform period.

Our decomposition exercise provides further insights on the sources of changes in the earnings gap. During the period of 1989–93, changes in observed characteristics—the "observed *X*'s effect"—accounted for more than half of the increase in the earnings gap, and changes in their prices—"the observed prices effect"—narrowed the gap by 1 percentage point. Among the observed characteristics that we have controlled for, the most noticeable change was the decline in women's relative experience, which was largely driven by the fact that women left employment much earlier than men. Figure 1 shows that while women's probability of leaving employment increases considerably after age 40, it only starts to increase after age 50 for men. Because older workers tend to earn more than younger workers, a higher job separation rate of older women widens the gender earnings gap. The observed prices effect is mainly attributable to the declining relative earnings of SOU employees. This is due to the fact that earnings in the private sector grow at a much faster pace than those in the state-owned sector. Because men are more likely to work in the state sector, a decline in the SOU premium reduces the gender earnings gap. Increase in the prices of unobserved skill—the "unobserved prices effect"—accounted for another half of the increase in the earnings gap, suggesting that rising earnings inequality adversely affected women's relative earnings.

During the period 1993–2004, changes in observed characteristics narrowed the earnings gap while increases in earnings inequality adversely affected women's relative earnings. The former was mainly driven by improvements in the relative education level of employed women. However, we do not observe any significant improvements in women's relative education level over the same period once nonemployed adults are included in the comparison. For example, while employed women's average education level was about 0.3 years lower than that of employed men in 1997, they were better educated than employed men in 2004. In contrast to the improvement in employed women's relative education level, the gender difference in education level for the entire population stayed at around 1.3 years throughout the entire sample period. This contrast suggests that women's employment status was more sensitive to education than that of men. Similar results have also been documented by Giles et al. (2006). In addition to the improvements in observed skills, we also observe improvements in unobserved skills. The improvements in unobserved skills of employed women reduced the earnings gap by 1.3 percentage points between 1993 and 1997 and another 2.5 percentage points between 1997 and 2004. Because women's employment rate declined considerably over the 1993–2004 period, particularly for less educated women, we do not know whether these improvements reflect genuine improvements in women's unobserved skills or a stronger positive selection on unobserved skills.

Consistent with the decomposition results, our analyses on employment retention rate also suggest that better-educated workers and those with a higher level of unobserved skills have a higher probability of being employed over two consecutive surveys in the later reform period. This type of positive selection plays a more significant role in women's employment. Age is another important determinant of employment, as retirement was reported to be the major reason for leaving employment for both men and women throughout the entire sample period. The population share of retirees grew steadily over time. Given that we have restricted our sample to 16–60-year-olds, the above evidence suggests that more and more people decided (or were forced) to retire before reaching the official mandatory retirement age. For women who were not working and not retirees in the survey year, more and more of them chose to stay at home as household wives. Actually, in the 2000 and 2004 surveys, the number of women who chose to be household wives was larger than those who were actively looking for jobs. All these pieces of evidence suggest that the recent market-originated reforms have a stronger negative effect on the employment of urban women.

The paper is structured as follows. The next section provides some background knowledge about the process of economic reforms in urban China and a brief discussion of the data source. Section III documents the evolution of the gender earnings gap between 1989 and 2004. Section IV analyzes the determinants of employment rate. A short conclusion is given in Section V.

II. Institutional Background and the Data

Before the economic reform, both employment and compensation in China were controlled by the state. Employers had little control over whom they could employ and how much to pay. The first step taken by the government to increase labor market flexibility is the reintroduction of bonuses and piece wages in 1978. However, rather than using bonuses as an incentive mechanism, most employers distributed bonuses equally among their employees so as to minimize contention (Walder 1987). The introduction of a labor contract system in the mid-1980s also failed to adding more flexibility to the Chinese labor market as contract workers were largely treated the same as permanent workers. Arguably, the most influential urban reform, known as *xia gang*, was introduced in the mid-1990s when the state-owned sector started to lay off workers on a large scale. *Xia gang* was first given a trial in 1994 and finally launched in 1997 (Appleton et al. 2002). As a result of the mass layoffs, the employment shares of SOEs and COEs have decreased considerably since 1997. Table 1 shows that both the total employment and the number of employees in SOUs increased year by year between 1984 and 1995. The employment share of SOUs stayed at around 73% over this period with little year-to-year variation. After that, the total employment and SOU employment started to fall, with the latter outpacing the former. As a result, SOU's employment share decreased from 73.5% in 1995 to only 60.9% in 2004. During the same period, the employment share of other-ownership units increased from 5.9% to 31.1%.

The mass layoff had a larger negative impact on women. Figure 1 plots men's and women's employment rates by age in 1989, 1993, and 2004. While

TABLE 1
TOTAL NUMBER OF "STAFF AND WORKERS," BY REGISTRATION TYPE

		Number of Workers in 10,000			Employment Share in %		
Year	Total	State Owned	Collective Owned	Other	State Owned	Collective Owned	Other
1984	11,890	8,637	3,216	37	72.6	27.0	.3
1985	12,358	8,990	3,324	44	72.7	26.9	.4
1986	12,809	9,333	3,421	55	72.9	26.7	.4
1987	13,214	9,654	3,488	72	73.1	26.4	.5
1988	13,608	9,984	3,527	97	73.4	25.9	.7
1989	13,742	10,108	3,502	132	73.5	25.5	1.0
1990	14,059	10,346	3,549	164	73.6	25.2	1.2
1991	14,508	10,664	3,628	216	73.5	25.0	1.5
1992	14,792	10,889	3,621	282	73.6	24.5	1.9
1993	14,849	10,920	3,393	536	73.5	22.9	3.6
1994	14,849	10,890	3,211	747	73.3	21.6	5.0
1995	14,908	10,955	3,076	877	73.5	20.6	5.9
1996	14,845	10,949	2,954	942	73.8	19.9	6.3
1997	14,668	10,766	2,817	1,085	73.4	19.2	7.4
1998	12,337	8,809	1,900	1,628	71.4	15.4	13.2
1999	11,773	8,336	1,652	1,785	70.8	14.0	15.2
2000	11,259	7,878	1,447	1,935	70.0	12.9	17.2
2001	10,792	7,409	1,241	2,142	68.7	11.5	19.8
2002	10,558	6,924	1,071	2,563	65.6	10.1	24.3
2003	10,492	6,621	951	2,920	63.1	9.1	27.8
2004	10,576	6,438	851	3,287	60.9	8.0	31.1

Source. The information is extracted from table 1-14 of the China Labour Statistical Yearbook (2005).

the age-employment profile did not change much between 1989 and 1993, it shifted downward considerably in 2004. The magnitude of the shift was much more pronounced for women, particularly for women who were older than 45. Women's higher probability of retiring early and higher risk of being laid off were both accountable for the larger decline in women's employment rate. Appleton et al. (2002) document that the incident rate of layoff was 12% for men and 22% for women. After losing their jobs, many people left the labor market. Giles et al. (2006) show that the labor force participation rate dropped from 74.4% in January 1996 to 63.1% in November 2001 for women and from 93.0% to 86.3% for men. Conditional on being laid off, women generally have to face a higher unemployment rate. In November 2001, the unemployment rate of the 40–50-year-olds was 10.3% for men and 17.1% for women. Consequently, they also had a lower reemployment probability. For example, Giles et al. (2006) show that while 44.3% of 40–50-year-old men were reemployed within 12 months of leaving their jobs, the corresponding figure for women was only 22.1%. All those pieces of evidence suggest that it is important to jointly examine gender difference in earnings and employment.

In this study, we use data extracted from the 1989–2004 China Health and Nutrition Survey (CHNS). The CHNS is a longitudinal survey, designed jointly by the Carolina Population Center at the University of North Carolina at Chapel Hill and the Chinese Academy of Preventive Medicine. It is implemented by related city/county anti-epidemic stations under the organization of the Food Inspection Services of the provinces. Although the CHNS is designed to examine the effects of health, nutrition, and family planning policies, it does contain detailed income and job-related information that can be used to analyze labor market issues. The survey took place over a 3-day period and drew a sample of about 4,400 households with a total of 16,000 individuals in nine provinces, including Liao Ning, Hei Long Jiang, Jiang Su, Shan Song, He Nan, Hu Bei, Hu Nan, Guang Xi, and Gui Zhou. Households from Hei Long Jiang were only surveyed after the 1993 survey, and households from Liao Ning were not surveyed in 1997.

Within each province, its urban sample covers the provincial capital city and a lower income city when feasible. In two provinces, the survey has to use other larger cities rather than the provincial capitals. The average education level of the CHNS provinces is lower than that of the national average. Using the 1990 census, we find that 45% of the urban workers (aged 16–60) have at least a high school diploma at the national level, the corresponding value of the CHNS provinces is 44%, and the corresponding value of the CHNS data is only 11%. Therefore, we cannot apply our findings to the national level. Nevertheless, our results provide a useful complement to existing studies that tend to depend on samples overrepresented by more developed provinces. For example, 51.5% of Zhang et al.'s (2007) sample have at least a high school education.

The CHNS has a respectable retention rate. For example, among the 15,923 individuals who were surveyed in 1989, 14,028 were also surveyed in 1991. Even after 15 years of the initial survey, the CHNS still managed to survey 56.3% of the original sample population whose community still participated in the 2004 survey. The retention rate is even higher among those who were born between 1929 and 1969 (aged 20–60 in 1989), with a value of 61.4%. Since 1993, all new households formed from sample households residing in sample areas were added to the survey, resulting in a total of 13,893 individuals. Since 1997, additional households were added to replace those no longer participating.

In this study, we focus on individuals who were between 16 and 60 years old in the survey year, were not enrolled in any types of schools, and lived in urban areas.[2] Their earnings are measured by self-reported average monthly

[2] The main reason for restricting the sample to 16–60-year-olds is to make our sample comparable

wage plus bonus and subsidies. Due to the number of digits allocated for data recording, the monthly wage is capped at 999 yuan in 1991, at 9,999 yuan between 1993 and 2000, and 99,999 in 2004. Because the 1989 wages are constructed by the survey center using information on daily wages or piece rate, they are not capped in 1989. As a result, we have four observations whose monthly wages were greater than or equal to 999 yuan in 1989. It is interesting that nobody's earnings were affected by the cap between 1991 and 2004. However, in 2004, four workers' wages were greater than or equal to 9,999 yuan, which would be capped at 9,999 if the cap was not increased in 2004. For consistency, these eight observations are excluded.[3] Because most self-employed or family workers did not report their monthly earnings, they are excluded from our earnings analyses. However, they are treated as employed and included in our employment analyses. The nominal wages are deflated by the CPI (1978 = 100) of all urban areas.

Another labor market variable used in our analysis is the type of registration of an employee's work unit. Although the survey question on this issue varies over time, two groups of workers can be consistently identified throughout the entire sample period: those who worked for SOUs and those who worked for COEs. Consequently, we group all employed workers into three categories: SOU, COE, and others. The last group largely consists of employees of private enterprises, foreign-owned enterprises, overseas Chinese-owned enterprises, and joint ventures. In later discussion, we will refer this group as private employees.

Other variables used in the analysis are the province of residence, years of schooling, potential years of experience, and relative height. Years of schooling is constructed from two variables: years of formal education completed and the highest level of education obtained. The existence of two education-related variables enables us to check the consistency of the education information. Presumably, the reported years of schooling is likely to be accurate if these two measures conform with each other. Therefore, reported years of schooling is used whenever these two variables are consistent. If these two variables are only consistent in some survey years, then the years of schooling in these years are used for the nearest inconsistent years. If these two variables are inconsistent throughout the entire sample period, then the respondent's average years of schooling is used for each period. Average years of schooling over the entire

to the ones used in previous studies. One potential problem of including workers who are older than 55 is that women's mandatory retirement age is 55, which is younger than the selected up-bound. To see whether our results are sensitive to the age restriction, we tried to limit our analysis to the 16–45-year-olds. While the level of the gender gap of the restricted sample is slightly lower than that of the entire sample, its trend does not depend on the age restriction.

[3] Among these eight excluded individuals, six of them are males.

sample period might differ from actual schooling in a particular year for individuals whose formal education consists of several disconnected intervals. In our data, only 64 individuals went back to school after leaving school in the previous survey, and only 17 individuals' education variables were not consistent with each other. Therefore, we are confident that the benefits of using our constructed years of schooling rather than the reported years of schooling outweigh the costs.

Potential years of working experience are calculated as age minus years of schooling minus 6. It should be noted that measurement errors introduced by using the constructed series as a proxy for actual experience may differ across cohorts and provinces. This is because China's education system has experienced considerable changes between 1960 and 1980, and the timing of these changes differs across geographic areas. For example, while it normally takes 12 years to complete high school, it could take as little as 9 years to complete high school in some cities in the 1970s. To check whether age is a better proxy for experience, we tried to use age and age squared instead of potential experience and its square in the earnings regression. The estimated gender gap is not sensitive to which measure is used. However, the rate of return to education is considerably lower when age is used as a proxy for experience. To keep our results comparable with previous studies, we only report the estimation results that use potential years of experience.

We include relative height, defined as the ratio of deviation from a gender-specific mean to the standard deviation of the corresponding gender's height distribution, to capture the potential influence of family background. Many researchers (e.g., Persico, Postlewaite, and Silverman 2004; Case and Paxson 2008) have tried to identify and explain the height premium. One explanation is that height partially captures the impact of family background. As long as the lack of nutrition is a concern for child development, we should expect that taller people are more likely to come from affluent families. Because of the positive correlation between earnings of parents and of their children, people raised by affluent families tend to earn more than others. We indeed find that taller workers earn more using the CHNS data.

Table 2 reports the basic sample statistics by gender and survey year. The female to male earnings ratio fell from 89.5% in 1989 to 82.0% in 1997, which is consistent with the evidence documented by previous studies. Interestingly, women's relative earnings gained some ground after that. Due to the large standard errors, year-to-year changes in earnings ratio are not statistically significant even at the 10% level. One noticeable pattern emerging from table 2 is that the narrowing earnings gap in the 1997–2004 period was accompanied by a decline in employment rate which was steeper for

TABLE 2
DESCRIPTIVE STATISTICS

	1989	1991	1993	1997	2000	2004
Earnings ratio	.895	.832	.826	.820	.847	.893
A. Males:						
Entire sample:						
Employment rate	.922	.926	.903	.881	.801	.695
No. of earners	.901	.783	.793	.716	.625	.395
Schooling	8.10	8.35	8.77	9.03	9.60	9.80
Age	35.10	36.22	36.89	37.57	39.52	41.55
Employed:						
Schooling	9.08	9.48	9.80	10.23	10.94	11.13
Age	36.05	36.09	37.36	37.74	39.33	40.98
Monthly earnings	54.24	70.80	94.20	111.65	171.76	220.49
SOU employees	.705	.718	.704	.630	.611	.607
Collectives	.262	.252	.247	.271	.225	.116
Daily working hours	7.94	7.98	7.83	7.91	7.92	8.02
B. Females:						
Entire sample:						
Employment rate	.836	.806	.790	.758	.683	.526
No. of earners	1.001	.902	.793	.762	.646	.464
Schooling	6.84	6.98	7.29	7.71	8.52	8.65
Age	34.84	36.10	37.27	37.78	39.29	42.08
Employed:						
Schooling	8.68	9.21	9.29	9.96	10.83	11.30
Age	33.58	32.99	34.43	34.75	36.71	37.78
Monthly earnings	48.54	58.88	77.80	91.54	145.49	196.80
SOU employees	.601	.638	.635	.597	.571	.566
Collectives	.350	.334	.328	.275	.252	.125
Daily working hours	7.87	7.95	7.85	7.79	7.85	7.84

Note. The sample consists of individuals aged between 16 and 60. No. of earners refers to the number of employed adults in the household excluding the respondent.

women. For instance, while men's employment rate declined by 18.6 percentage points between 1997 and 2004, women's employment rate decreased by 23.2 percentage points. This finding is consistent with the statistics documented by the National Bureau of Statistics (2001) using the second survey on women's status in China. According to the National Bureau of Statistics (2001), the urban employment rate of 18–64-year-olds fell from 90.0% in 1990 to 81.5% in 2000 for men, and from 76.3% to 63.7% for women.

To see whether the characteristics of employed individuals differ from those of the entire population, in table 2 we separately report years of schooling and age for the entire sample and the employed sample. Women's average years of schooling was about 1 year lower than that of men, but the difference was much smaller between employed men and women. Actually, in the last survey, employed women were better educated than employed men. Zhang et al. (2007) also find that employed women were better educated than employed men after 2001. The narrowing education gap suggests that less educated women exited from employment at a higher rate than less educated men.

Because employed women were less educated than employed men in the early period, even if education had the same impact on men's and women's employment, a rise in education requirement would have a larger negative impact on women's employment. However, a gender-neutral change can narrow the education gap but cannot reverse it. Therefore, the changes in the education gap documented in table 2 suggest that women's employment status was more sensitive to schooling. A similar pattern has also been found by Giles et al. (2006). The growing impact of education on employment narrowed the gender earnings gap. A similar conclusion has been drawn by Hunt (2002) using a German data set.

While schooling has a positive effect on employment, age seems to have a negative effect on women's employment. On average, employed women were about 1 year younger than the entire female population at the start of the sample period and around 4 years younger at the end of the sample period. The increase in the age difference between employed women and the entire sample indicates that the difference between the employment rate of older and younger women has increased over time. Overall, the statistics reported in table 2 show that both the earnings gap and employment gap changed during the 1989–2004 period. To have a deeper understanding of the impact of economic reforms on the labor market, we need to jointly examine changes in these two variables.

III. Wage Results

A. Mean Regression Results

To depict a general picture of the evolution of the gender earnings gap, we first use the CHNS as six cross-section data sets. We run log earnings regressions for each of the six waves separately. The control variables include a gender dummy (= 1 for male and 0 otherwise), years of education, potential years of experience and its square, relative height, the registration type of work unit, and the province of residence. To help us understand the contributions of these observed variables to the earnings gap, we add them into our regression gradually. The regression results are reported in table 3.

In panel A of table 3, a gender dummy is used as the only control variable. Similar to table 2, the largest increase in the gender earnings gap occurred between 1989 and 1991. The gap reached its peak in 1997 and has declined slightly since then. Panel B shows that controlling for education and experience reduces the gap by 1.9 to 4.5 percentage points.[4] Including types of work

[4] Hereafter, the coefficient on gender dummy γ will all be converted to percent using the formula $(\exp(\gamma) - 1)$, as suggested by Halvorsen and Palmquist (1980).

TABLE 3
ESTIMATION RESULTS FROM EARNINGS REGRESSIONS

	1989	1991	1993	1997	2000	2004
A. Gender	.1284	.1724	.1798	.1837	.1746	.1722
	(.0254)***	(.0188)***	(.0267)***	(.0259)***	(.0350)***	(.0385)***
B. Gender	.1095	.1271	.1543	.1550	.1532	.1488
	(.0248)***	(.0173)***	(.0262)***	(.0253)***	(.0341)***	(.0358)***
Schooling	.0235	.0256	.0155	.0353	.0507	.0823
	(.0046)***	(.0028)***	(.0043)***	(.0044)***	(.0059)***	(.0066)***
Experience	.0178	.0185	.0271	.0177	.0182	.0129
	(.0059)***	(.0026)***	(.0040)***	(.0041)***	(.0057)***	(.0062)**
Experience^2/100	−.0112	−.0111	−.0405	−.0204	−.0192	−.0069
	(.0165)	(.0057)*	(.0084)***	(.0092)**	(.0126)	(.0137)
Relative height	.0326	.0148	.0157	.0369	.0434	.0496
	(.0141)**	(.0095)	(.0142)	(.0138)***	(.0191)**	(.0174)***
C. Gender	.1073	.1276	.1543	.1579	.1507	.1464
	(.0248)***	(.0173)***	(.0259)***	(.0250)***	(.0339)***	(.0357)***
Schooling	.0214	.0271	.0218	.0325	.0425	.0781
	(.0047)***	(.0029)***	(.0045)***	(.0045)***	(.0063)***	(.0070)***
Experience	.0169	.0188	.0293	.0192	.0157	.0131
	(.0060)***	(.0026)***	(.0039)***	(.0041)***	(.0058)***	(.0062)**
Experience^2/100	−.0093	−.0113	−.0431	−.0231	−.0167	−.0077
	(.0166)	(.0057)**	(.0083)***	(.0091)**	(.0126)	(.0136)
Relative height	.0329	.0151	.0157	.0389	.0403	.0515
	(.0141)**	(.0096)	(.0140)	(.0137)***	(.0190)**	(.0175)***
SOU employees	.0932	−.0892	−.2972	−.1222	.1253	.0207
	(.0744)	(.0556)	(.0644)***	(.0424)***	(.0507)**	(.0420)
Collectives	.0565	−.0627	−.1798	−.2336	−.0187	−.1355
	(.0752)	(.0564)	(.0666)***	(.0452)***	(.0552)	(.0606)**
D. Gender	.1054	.1287	.1475	.1563	.1490	.1553
	(.0243)***	(.0167)***	(.0253)***	(.0242)***	(.0330)***	(.0336)***
Schooling	.0201	.0293	.0262	.0321	.0423	.0776
	(.0047)***	(.0029)***	(.0045)***	(.0045)***	(.0063)***	(.0068)***
Experience	.0198	.0187	.0270	.0214	.0188	.0207
	(.0059)***	(.0025)***	(.0039)***	(.0040)***	(.0057)***	(.0059)***
Experience^2/100	−.0205	−.0103	−.0380	−.0296	−.0242	−.0258
	(.0164)	(.0056)*	(.0082)***	(.0090)***	(.0125)*	(.0130)**
Relative height	.0169	.0245	.0258	.0132	.0115	.0447
	(.0152)	(.0103)**	(.0155)*	(.0144)	(.0201)	(.0174)**
SOU employees	.0736	−.1001	−.2647	−.1067	.1659	.0721
	(.0729)	(.0538)*	(.0636)***	(.0416)**	(.0507)***	(.0415)*
Collectives	.0111	−.0983	−.2013	−.2553	−.0111	−.1336
	(.0740)	(.0545)*	(.0652)***	(.0444)***	(.0546)	(.0581)**
Number of observations	949	1,320	1,136	1,230	1,095	879

Note. The sample consists of individuals aged between 16 and 60. Gender = 1 for males. In panel D, the residence of province is also used as a control variable. The log of real monthly average earnings is used as the dependent variable. Numbers in parentheses are standard errors. The standard errors are corrected for the potential correlation among individuals of the sample household.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

units in panel C and the province of residence in panel D has no obvious impact on the estimated earnings gap. After including all control variables, the gender earnings gap measured by the coefficient on the gender dummy increased from 0.1054 log points (11.1%) in 1989 to 0.1287 (13.7%) in 1991 and then to 0.1563 (16.9%) in 1997. These estimates are very close to Yang (2005), who finds that the coefficient on the gender dummy is 0.097 in 1988 and 0.155 in 1995. Interestingly, the gender earnings gap stabilized around 0.15 until the end of the sample period, which differs from Zhang et al. (2007), who show that the gap increased by 8 percentage points from 2001 to 2004 before controlling for the observed characteristics.

The difference between Zhang et al. (2007) and the current study is likely due to differences in the sample coverage. Zhang et al.'s (2007) sample covers more developed provinces than the CHNS sample. These more developed provinces generally have a more viable private sector that can absorb more low-skilled workers. Hence, the employment prospects of low-skilled workers in these provinces were better than those of their counterparts in less developed provinces. As a result, the low-skilled workers' employment share is larger in Zhang et al. (2007). This argument is supported by the fact that while Zhang et al. (2007) show that the relative education level of employed women declined from 2001 to 2004, our calculation using the CHNS data shows the opposite. A decline in relative skill level of women raises the gender earnings gap. The difference between Zhang et al. (2007) and the current study suggests that the evolution of the gender earnings gap might follow different paths across provinces and could depend on the level of economic development. Therefore, we should be cautious when trying to extrapolate findings from any particular data set to the national level.

Compared to the considerable increase in the gender earnings gap between 1989 and 1997, the rate of return to education (as reported in panel D) was very stable during that period. It only increased from 0.020 in 1989 to 0.032 in 1997. The relatively stable rate of return to education in the 1989–97 period suggests that changes in the returns to education were not responsible for the early increase in the gender earnings gap. Interestingly, the rate of return to education increased considerably in the 1997–2004 period when the gender earnings gap was relatively stable. Our estimated rates of return to education for the period of 1989–97 are slightly lower than what have been documented by several previous studies. For instance, both Liu (1998) and Yang (2005) find that the rate of return to education was around 0.03 in 1989, and Yang (2005) finds that the rate of return increased to 0.059 in 1995. Similarly Zhang et al. (2007) find that rate of return to education increased from 0.02 in 1988 to 0.044 in 1994. Nevertheless, our 2004 estimate is comparable to that of Zhang et al. (2007). One potential reason for our low estimate is that average years of education is lower in the CHNS than in the CHIP used by Yang (2005) and the Urban Household Survey used by Zhang et al. (2007). For example, Yang (2005) shows that average years of schooling was 10.7 in 1988 and 11.8 in 1995. In contrast, the corresponding value of our data was 8.9 in 1989, 9.6 in 1993, and 10.1 in 1997. If the rate of return to education is mainly driven by the reward of obtaining a high school diploma or a bachelor's degree, then the estimate will be lower if a sample is overrepresented by less educated workers. If we reweight the 1997 CHNS sample so that its education composition is comparable to that of the CHIP data and run the wage regression using the reweighted data, the estimated rate of return to education becomes 4.4, which is the same as the one reported by Zhang et al. (2007).

Urban reforms also had a profound effect on earnings differentials among SOU, COE, and the private sector. In 1989, the earnings of SOU employees and of COE employees did not significantly differ from those of private employees, but the earnings of SOU employees were significantly higher than those of COE employees. As urban economic reforms kicked in during the early 1990s, both SOU employees and COE employees gradually lost out to private employees. In 1993, SOU workers earned about 23% less than private employees and COE employees earned about 18% less. After that, the earnings position of SOU workers improved steadily, and they earned more than private employees after 2000. However, COE employees always earned less than workers in other sectors. Because men have a higher probability of working in SOUs and a lower probability of working in COEs than women, changes in the relative earnings position of SOU workers will affect the gender earnings gap.

So far we have not discussed the potential impact of changes in working hours on the gender earnings gap. This is mainly because we cannot construct a consistent monthly working hours series from the data. If changes in working hours were systematically different across genders, then failing to control for working hours biases our results. We can partially address this issue by controlling for daily working hours, which is consistently recorded in the data. Table 4 reports the estimation results where daily working hours are also included as a control variable. For the sake of brevity, we only report the estimation results with the full set of control variables. Interestingly, controlling for daily working hours only slightly reduces the coefficient on the gender dummy. This is because women's daily working hours are only about 0.1 hours shorter than that of men. Moreover, adding daily working hours does not affect the trend of the gender earnings gap.

TABLE 4
ESTIMATION RESULTS AFTER CONTROLLING FOR DAILY WORKING HOURS

	1989 (1)	1991 (2)	1993 (3)	1997 (4)	2000 (5)	2004 (6)
Gender	.1042	.1269	.1455	.1569	.1471	.1452
	(.0244)***	(.0166)***	(.0252)***	(.0243)***	(.0335)***	(.0333)***
Schooling	.0198	.0292	.0277	.0323	.0420	.0779
	(.0047)***	(.0028)***	(.0045)***	(.0045)***	(.0064)***	(.0067)***
Experience	.0201	.0184	.0261	.0217	.0193	.0226
	(.0059)***	(.0025)***	(.0039)***	(.0040)***	(.0058)***	(.0059)***
Experience^2/100	−.0216	−.0099	−.0338	−.0301	−.0233	−.0268
	(.0165)	(.0055)*	(.0085)***	(.0090)***	(.0126)*	(.0130)**
Relative height	.0178	.0236	.0248	.0127	.0061	.0544
	(.0154)	(.0102)**	(.0154)	(.0144)	(.0206)	(.0172)***
Daily working hours	.0062	.0097	.0362	.0183	−.0070	.0215
	(.0145)	(.0128)	(.0161)**	(.0127)	(.0154)	(.0136)
SOU employees	.0688	−.0947	−.2738	−.0848	.1381	.0785
	(.0730)	(.0539)*	(.0662)***	(.0422)**	(.0514)***	(.0416)*
Collectives	.0075	−.0922	−.2118	−.2382	−.0114	−.1314
	(.0741)	(.0547)*	(.0676)***	(.0450)***	(.0556)	(.0581)**
Number of observations	941	1,307	1,122	1,211	1,053	854

Note. The sample consists of individuals aged between 16 and 60. Gender = 1 for males. The log of real monthly average earnings is used as the dependent variable. The residence of province is also used as a control variable. Numbers in parentheses are standard errors. The standard errors are corrected for the potential correlation among individuals of the sample household.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

B. Quantile Regression Results

While the mean regression results reveal the gender earnings gap between an "average man" and an "average woman," they cannot tell how the gap varies across the earnings distribution. To examine the gender earnings gap at different points of the earnings distribution, we run a series of quantile regressions at the 10th, 25th, 50th, 75th, and 90th percentiles with the full set of controls, as in panel D of table 3. Unlike the OLS estimates that uncover the returns to observed characteristics at the sample mean, the estimates of quantile regressions reveal the returns to observed characteristics at the selected percentiles.[5] The coefficients on the gender dummy are plotted in figures 2 and 3.

Except for the 2004 results, the gender earnings gap at the top decile is larger than at the bottom decile and the gap at the top quartile is larger than that at the bottom quartile. However, the opposite is true in 2004. These changes show that the gender earnings gap evolved differently at different percentiles. It grew steadily at the 10th and 25th percentiles throughout the 1989–2004 period but only widened between the 1989–1993 period at the 75th and 90th percentiles. During the period 2000–2004, it experienced a 9

[5] The estimation results are reported in tables A1–A5.

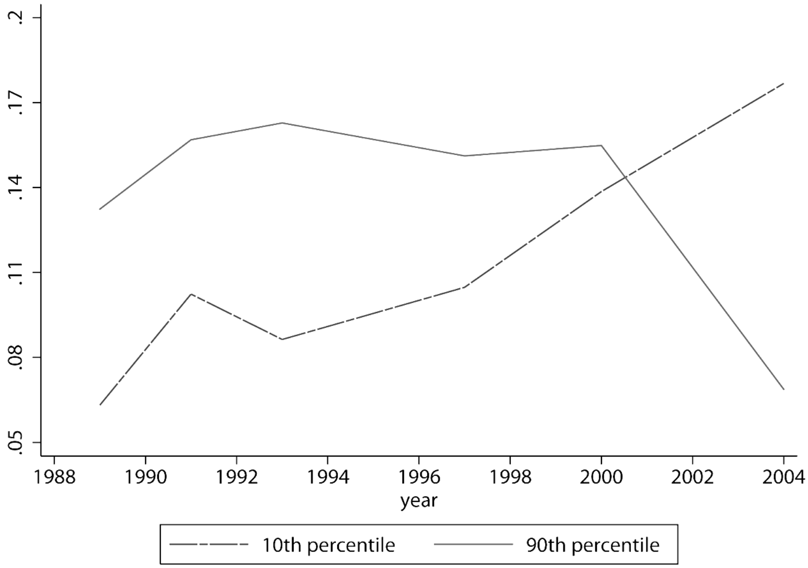

Figure 2. The gender earnings gap at the 10th and 90th percentiles

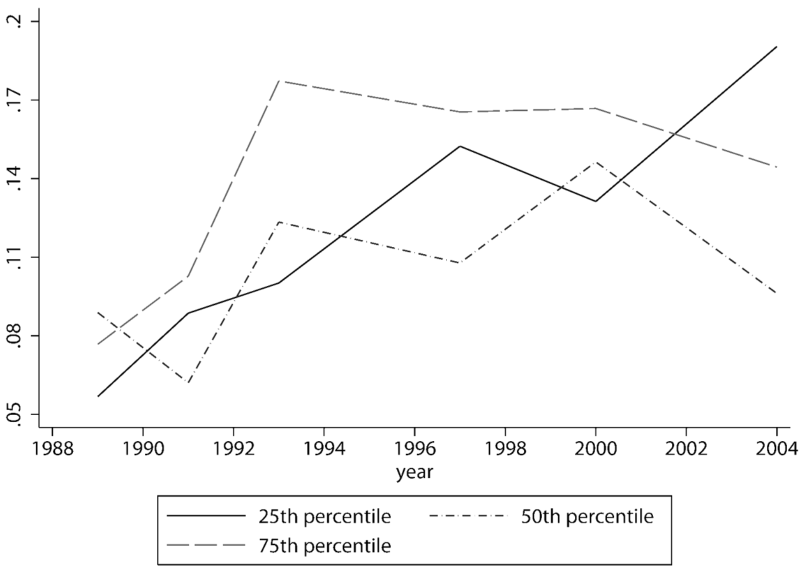

Figure 3. The gender earnings gap at the 25th, 50th, and 75th percentiles

percentage point drop at the 90th percentile and a 5 percentage point drop at the 75th percentile. These estimation results suggest that the relative stability in the gender earnings gap at the mean is the result of two opposing forces: one widened the gap at the bottom of the earnings distribution and another one narrowed the gap at the top of the earnings distribution. Similar to our results, Zhang et al. (2007) also find that the magnitude of the increase in the earnings gap between the 1989–2004 period is much larger at the bottom decile than at the top decile of the earnings distribution.

One potential explanation for the steady deterioration of female workers at the bottom of the earnings distribution is that lower-paid women are more likely to work in sectors with relatively slow earnings growth. Our calculation shows that while the average earnings increased by 2.27 times between 1989 and 2004 in the service sector where women accounted for about 70% of its workforce, the earnings of professionals increased by 2.94 times, of whom 60% were men. If women at the lower end of the wage distribution have a higher probability of working at jobs with slower earnings growth, then the earnings gap at the lower end will widen over time. Another potential explanation is the increase in the price of unobserved skill. As long as women's level of unobserved skills is lower than that of men, a rising price will widen the gender earnings gap at both the bottom and the top of the earnings distribution. Unfortunately, it is not clear to us why the gender gap narrowed from 2000 to 2004 at the top of the earnings distribution.

C. Decomposition Results

To spell out the contributions of various factors to the gender earnings gap, we adopt the decomposition method proposed by Blau and Kahn (1997). The main advantage of Blau and Kahn (1997) decomposition over the traditional Oaxaca-Blinder decomposition is that the former can identify the contribution of earnings inequality to the gender earnings gap.[6] The idea of decomposing changes in earnings inequality into a price factor and a skill factor was originated by Junh, Murphy, and Pierce (1991). If the residuals from a wage regression can be viewed as a payment to unobserved skills, then we can use a worker's percentile ranking in the residual wage distribution as a measure for his unobserved skills, and the dispersion of the residual as the "price" of the unobserved skill. The gender earnings gap, D_t, can be modeled as

$$D_t = \overline{\ln w_{mt}} - \overline{\ln w_{ft}} = \Delta X_t \beta_t + \sigma_t \Delta \theta_t, \tag{1}$$

where $w_j (j = m, f)$ represents monthly earnings, and X is a vector of control

[6] For more detailed discussion on the advantages of the Blau and Kahn (1997) decomposition, please refer to Altonji and Blank (1999).

variables including education, experience, relative height, registration type of work units, and residence of province; the subscript m means male and f female; $\overline{X}$ represents the sample average of X; β is the price of the corresponding personal characteristics; θ is the standardized residual; σ is the standard deviation of the earnings residuals; and Δ denotes the male-female difference in X and θ at their corresponding means. If we believe that men are paid fairly according to their personal characteristics, then the coefficients from men's earnings regression measure the prices of the corresponding characteristics. Therefore, we can use $\hat{\beta}_{mt}$ for β_t and $\hat{\sigma}_{mt}$ for σ_t in the decomposition exercise.

The change in the earnings gap between year t and 0 can then be decomposed as:

$$\begin{aligned} D_t - D_0 &= (\Delta X_t - \Delta X_0)\beta_t + \Delta X_0(\beta_t - \beta_0) \\ &\quad + (\Delta\theta_t - \Delta\theta_0)\sigma_t + \Delta\theta_0(\sigma_t - \sigma_0). \end{aligned} \tag{2}$$

The first term reflects the contribution of changes in observed characteristics to the earnings gap when evaluated at period t price. The second term reveals the contribution of changes in returns to characteristics evaluated at period 0 difference in characteristics. These two terms are called "observed X's effect" and "observed prices effect," respectively, by Blau and Kahn (1997), and their sum corresponds to the "explained" component of the traditional Oaxaca-Blinder decomposition. The third term represents the contribution of changes in the percentile ranking of the female earnings residuals while holding the standard deviation of male earnings residuals at period t level, called the "gap effect." Finally, the fourth term of equation (2), the "unobserved prices effect," measures the contribution of changes in the standard deviation of male earnings residuals holding the percentile rankings of female earnings residuals at their period 0 levels. The sum of the last two terms is the "unexplained" component of the traditional Oaxaca-Blinder decomposition.

The last two terms of equation (2) can be rewritten as

$$\begin{aligned} &(\Delta\theta_t - \Delta\theta_0)\sigma_t + \Delta\theta_0(\sigma_t - \sigma_0) = \\ &[(\theta_{tm} - \theta_{tf}) - (\theta_{0m} - \theta_{0f})]\sigma_{tm} + (\theta_{0m} - \theta_{0f})(\sigma_{tm} - \sigma_{0m}). \end{aligned} \tag{3}$$

Because the averages of both θ_{tm} and θ_{0m} are zero, expression (3) can be further simplified as

$$(\Delta\theta_t - \Delta\theta_0)\sigma_t + \Delta\theta_0(\sigma_t - \sigma_0) = (-\theta_{tf} + \theta_{0f})\sigma_{tm} - \theta_{0f}(\sigma_{tm} - \sigma_{0m}), \tag{4}$$

where $\theta_{tf}\sigma_{tm}$ and $\theta_{0f}\sigma_{0m}$ are the averages of female earnings residuals in year t and 0 imputed based on the corresponding year's men's earnings regressions. To impute $\theta_{0f}\sigma_{tm}$, we first estimate each woman's percentile ranking in year

TABLE 5
DECOMPOSITION OF CHANGES IN THE GENDER EARNINGS GAP

	1989–93 (1)	1993–97 (2)	1997–2004 (3)	1989–2004 (4)
Changes in the earnings gap	.051	.004	−.011	.044
Observed *X*'s	.026	−.025	−.019	−.035
Schooling	−.002	−.005	−.026	−.048
Experience	.056	.003	.019	.070
Experience^2/100	−.042	−.002	−.010	−.046
Relative height	.001	.000	.003	.003
SOU employees	.004	.010	.002	−.003
Collectives	.004	−.029	.001	−.004
Provinces	.005	−.002	−.006	−.007
Observed prices	−.010	.015	.018	.041
Schooling	.000	.003	.012	.029
Experience	.006	−.016	.007	.005
Experience^2/100	−.000	.009	−.000	−.000
Relative height	−.000	.001	−.003	−.001
SOU employees	−.028	.006	.008	.005
Collectives	.013	.012	−.000	.003
Provinces	−.000	.002	−.006	−.000
Gap effect	.011	−.013	−.025	−.011
Unobserved prices	.024	.026	.014	.049
Mean female residual percentile:				
1989	40.36			40.36
1993	39.14	39.14		
1997		38.69	38.69	
2004			41.63	41.63

Note. The sample consists of individuals aged between 16 and 60. The prices of observed characteristics used in the decomposition are estimated using the male sample.

0 based on her earnings residual in that year's distribution of men's earnings residuals. Then, we assign her the residual of the corresponding percentile of the year t distribution of men's earnings residuals. For example, if a woman's year 0 earnings residual is ranked at the pth percentile in that year's men's earnings residual distribution, then her predicted year t earnings residual, $\theta_{0f}\sigma_{tm}$, will be set at the residual of the pth percentile of the year t men's earnings residual distribution.

Table 5 reports the decomposition results. The β's used in the decomposition are estimated using only men's earnings with the same control variables as those that were used in panel D of table 3. Given the relatively low employment rate in the 2000s, it is possible that the estimates from the men's earnings regression might be also subject to selection bias. Therefore, we also tried to estimate β's using the full maximum likelihood estimates of the Heckman selection model. We find that the decomposition results are not sensitive to how β's are estimated. We suggest that this is because the low male employment rate is mainly due to the difficulty of locating a job rather than to a low labor force participation rate. For simplicity, the discussion will focus on the results based on the OLS estimates.[7] To make our results comparable with existing studies, we group the entire sample into three periods: 1989–93, 1993–97, and 1997–2004. The impact of self-selection on women's earnings will be reflected in either the observed *X*'s effect or the gap effect.

The first column of table 5 reports the decomposition of changes in the earnings gap between 1989 and 1993, a period when the earnings gap increased the most. Changes in observed *X*'s widened the earnings gap by 2.6 percentage points. The decline in women's average years of experience was the major contributor. In 1989, the average years of experience of women was 2.06 years lower than that of men. The difference was 2.42 in 1993. The drop in women's average years of experience was mainly because older women had a higher probability of leaving employment than older men. Changes in the prices of observed *X*'s narrowed the gap by 1 percentage point, which was mostly attributable to the decline in the relative earnings of SOU workers. Because males were more likely to work for SOUs, the decline in the relative earnings had a larger impact on their average earnings.

The gap effect widened the earnings gap by 1.1 percentage points and the unobserved prices effect enlarged the gap by another 2.4 percentage points. The former is in contrast to the German evidence documented by Hunt (2002), who finds that the gap effect reduced the gender earnings gap in Germany. One potential explanation for this phenomenon is that skills accumulated during the planned economy period by women might be less valuable than those accumulated by men. If this is the case, then returns to women's observed human capital should be lower as well. To test this argument, we pool women and men together and interact the gender dummy with the full set of control variables. The coefficient on the interaction between education and gender dummy ($= 1$ for men) is not significantly different from 0 even at the 10% level in 1989, but is -0.019 and significant at the 1% level in 1993, suggesting that the rate of return to women's education is higher. A similar pattern has also be documented by previous studies (e.g., Gustafsson and Li 2000; Zhang et al. 2005). The evidence does not support the productivity difference interpretation.

The increase in the gap effect could also be the result of a rise in the degree of labor market discrimination against women or a relative deterioration in women's unobserved skills. The latter explanation would be correct if women with a higher level of unobserved skills have a higher job separation rate than other women. Since observed and unobserved human capital are likely to be

[7] The estimated β's using only men's earnings are reported in table A6. For comparison purposes, the estimated β's using the women only sample are also reported in the table.

positively correlated, this argument also implies that women with a higher level of observed human capital have a higher job separation rate. Our decomposition shows a mixed picture in this regard. On the one hand, women's relative education improved slightly during this period, which suggests that it was the low-skilled women who had a higher job separation rate. On the other hand, the decline in relative experience suggests that it was the more experienced women who had a higher exit rate. The conflicting evidence suggests that further research is needed to sort out reasons for the increase in the gap effect during the 1989–93 period. Nevertheless, even if the increase is indeed caused by a higher job separation rate for women with a higher level of unobserved skills, it is still not clear why the job separation rate and skill levels were only positively correlated for women.

Column 2 of table 5 shows that the gross earnings gap only widened by 0.4 percentage points between 1993 and 1997. The relatively stable gross earnings gap is the result of two opposing forces. Gender-specific factors, that is, the observed X's effect and the gap effect, narrowed the gap by 2.5 and 1.3 percentage points, respectively, while the observed prices effect and the unobserved prices effect widened the earnings gap by 1.5 and 2.6 percentage points. The observed X's effect was mainly driven by the fact that while the employment share of COEs increased for men, it decreased for women. If the decline was the result of women leaving COEs for better-paid jobs, it improves women's relative earnings. If it was the result of women leaving for nonemployment, then it worsened women's employment prospects. Our calculation supports the latter interpretation. Among the 140 men who were employed by COEs in 1993, 87% of them were still employed in 1997. However, among the 133 women who were employed by COEs in 1993, only 74% of them were employed in 1997. Among the 34 women who were not employed in 1997, 15 of them retired and 10 out of the 15 retirees were younger than 55, the mandatory retirement age for women; another six stayed at home as housewives; and only five of them were actively looking for jobs. The evidence suggests that earlier retirement is one of the most important factors for women's job separation. The fact that only a small proportion of women were actively looking for jobs after losing their previous ones indicates that the prospect of reemployment became tougher for women in the late 1990s.

Column 3 of table 5 shows that the gross earnings gap declined by 1.1 percentage points from 1997 to 2004. Again, the two gender-specific factors are the main driving forces for the narrowing gender earnings gap. The observed X's effect reduced the gap by 1.9 percentage points, and the gap effect narrowed it by another 2.5 percentage points. Among all the observed characteristics, the improvement in women's education is the single most important factor for the narrowing gap. If education was the only factor that had changed over the 1997–2004 period, we would have observed a 2.6 percentage point reduction in the gender earnings gap. Unfortunately, the improvement is not the result of women received more education but rather of the higher job separation rate among less educated women. For example, while 72.2% of the employed men with at most a lower secondary education in 1997 were still employed in 2004, the corresponding number was only 55.6% for women.

The observed X's effect was almost counterbalanced by the observed prices effect; the latter widened the gap by 1.8 percentage points. The considerable increase in the returns to education in the 2000s is the main contributor. It alone widened the gap by 1.2 percentage points. The unobserved prices effect enlarged the earnings gap by another 1.4 percentage points.

Column 4 reports the decomposition results for the entire sample period. The gross earnings gap in 2004 was 4.4 percentage points higher than its 1989 level. The increase was mainly driven by changes in the wage structure. Changes in the returns to various observed characteristics added 4.1 percentage points to the earnings gap. The increase in the rate of return to education alone accounted for more than 50% of the 4.1 points increase. It raised the earnings gap by 2.9 percentage points. However, any further increases in the rate of return to education in the future could be beneficial to women's relative earnings as their education level already overtook men's in 2004. Increases in the price of unobserved skills widened the gap by another 4.9 percentage points. Any further increase in the price of unobserved skills will cause further deterioration of women's relative earnings. This is because although the average level of women's unobserved skills improved slightly during the 1989–2004 period, it was still more than 8% lower than that of men. Therefore, if the price of unobserved skills keeps rising in the future, the rise will widen the earnings gap further. About half of the increases in the earnings gap caused by changes in the wage structure were balanced out by gender-specific factors, particularly by the increase in employed women's relative education level.

D. Sensitivity Analyses

Changes in the Sample Coverage

Although the CHNS has a respectable sample retention rate among communities that were surveyed consecutively, a considerable number of communities were only surveyed in part of the sample period. As a result, many households were not in the sample for some years. For example, all communities from Liao Ning province were not surveyed in 1997 while communities from Hei Long Jiang province have been added only since 1993. Moreover, some households dropped out, and some new households were added. If households did not drop out of the survey at random, or the gender gap differs across provinces, then changes in the sample coverage will affect our estimation results.

To check whether our results are sensitive to changes in the sample coverage, we rerun our estimation using households that were in the sample throughout the entire sample period. By doing so, we exclude all households from Liao Ning and Hei Long Jiang provinces. Because some household members in the selected households were not surveyed in some years and older household members were excluded from our analyses once they were older than 60, the size of the restricted sample still varies across surveys. The estimated gender earnings gap with and without controlling for personal and job-related characteristics is reported in panel A of table 5. A comparison between tables 3 and 5 suggests that the estimated gender gaps using the selected sample are slightly higher than the estimates using the entire sample. However, the year-to-year change in the gender gap follows a similar pattern, namely, the gender gap increased considerably in the early 1990s and stabilized in the early 2000s.

One noticeable difference between these two tables is that while the estimated gender gap widened between 1993 and 1997 in table 3, it narrowed in table 5. This difference suggests that the widening gender gap reported in table 3 might be due to changes in the sample composition. Since all households from Liao Ning were replaced by households from Hei Long Jiang, if the gender gap is larger in Hei Long Jiang, then changes in the sample coverage could introduce an upward bias to the estimated gender gap in 1997 reported in table 3. To address this issue, we rerun the 1993 and 1997 earnings regressions on all households except for those from Liao Ning and Hei Long Jiang. The estimated gender gap is 0.157 in 1993 and 0.146 in 1997, which suggests that changes in the sample coverage are indeed responsible for the increase in the estimated gender gap between 1993 and 1997 documented in table 3. However, even after controlling for changes in sample composition, our results still show that the gender earnings gap stabilized in the early 2000s.

The Potential Impact of Early Retirement

One issue that is largely ignored by previous studies on the Chinese gender earnings gap is the difference in mandatory retirement age between men and women. It is set at 60 for men and 55 for women. The difference in actual retirement age could be even larger in the later reform period as women were more likely to prefer or be forced to retire early. Women's higher probability of retiring early enlarges the gender earnings gap, as senior workers tend to be paid more than junior workers. To check whether our results are driven by the difference in retirement age between men and women, we rerun the earnings regression using a sample of workers aged between 16 and 45. Presumably, the employment status of these workers was not affected by early retirement decisions.

Panel B of table 6 reports the estimated gender gap using earnings data of the younger sample. Similar to what have been documented in table 3, the results from the younger sample also suggest that the gender gap increased between 1989 and 1997 and then stabilized in the 2000s. Except for the 2000 results, the magnitude of the gender gap from a relative young sample is very close to the estimates using the entire sample.

IV. Employment Results

In the previous section, we show that in both observed and unobserved skills employed women managed to narrow the gaps with employed men. To examine whether this was the result of the differences in the probability of exiting from employment across gender and skill levels, we create five panel data sets from the total six surveys. Each panel consists of two adjacent surveys, 1989–91, 1991–93, 1993–97, 1997–2000, and 2000–2004. The sample is restricted to those who reported positive earnings in the first survey and were interviewed in the second survey.[8] Figure 4 plots the employment rate in the second survey by workers' earnings quintiles in the first survey. Self-employed individuals are treated as employed in our analyses. The quintiles are calculated using the pooled earnings distribution.

Women's employment rate is almost universally lower than that of men, and the differences vary across earnings distribution and time. During the periods 1989–91 and 1991–93, men's earnings in the first year did not affect their employment probability in the second year. However, women at the top of the earnings distribution have a considerably lower employment probability than women at the bottom of the earnings distribution. As more women at the top of the earnings distribution left their jobs, the gender earnings gap widened. We suggest that the negative correlation between earnings and employment is the result of women's higher probability of retiring early. Because earnings largely reflected workers' seniority in the early reform period, women's higher probability of retiring early widened the gender earnings gap.

The 1993–97 data show a slightly positive correlation between earnings

[8] To have a complete picture of the impact of changes in employment rate on the gender earnings gap we should compare both the exit rate and the entrance rate. However, because only a very small number of individuals who did not work in the first period were employed in the second period, the difference in the entrance rate between men and women is unlikely to have any significant impact on the earnings gap. For the sake of conciseness, we focus our discussion on the exit rate.

TABLE 6
SENSITIVITY ANALYSES

	1989 (1)	1991 (2)	1993 (3)	1997 (4)	2000 (5)	2004 (6)
	A. Surveyed throughout the Entire Sample Period					
Gender gap	.1352	.1952	.2111	.1714	.1897	.1911
	(.0523)***	(.0359)***	(.0385)***	(.0409)***	(.0506)***	(.0697)***
Personal characteristics	No	No	No	No	No	No
Ownership of work units	No	No	No	No	No	No
Residence of province	No	No	No	No	No	No
Gender gap	.1179	.1536	.1863	.1436	.1643	.1411
	(.0522)**	(.0347)***	(.0379)***	(.0376)***	(.0471)***	(.0621)**
Personal characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Ownership of work units	Yes	Yes	Yes	Yes	Yes	Yes
Residence of province	Yes	Yes	Yes	Yes	Yes	Yes
Observations	321	449	464	463	352	300
	B. Younger than 45 in the Survey Year					
Gender gap	.1284	.1438	.1549	.1740	.1215	.1730
	(.0254)***	(.0202)***	(.0297)***	(.0288)***	(.0410)***	(.0467)***
Personal characteristics	No	No	No	No	No	No
Ownership of work units	No	No	No	No	No	No
Residence of province	No	No	No	No	No	No
Gender gap	.1062	.1289	.1379	.1587	.1017	.1685
	(.0244)***	(.0184)***	(.0279)***	(.0266)***	(.0392)***	(.0407)***
Personal characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Ownership of work units	Yes	Yes	Yes	Yes	Yes	Yes
Residence of province	Yes	Yes	Yes	Yes	Yes	Yes
Observations	945	1,096	908	976	789	588

Note. Personal characteristics include years of education, potential years of experience and its square, and normalized relative height. Ownership of work units consists of SOU dummy and COE dummy. Residence of province includes six residence of province dummies in panel A and eight residence of province dummies in panel B.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

and employment for men. However, women at the top of the earnings distribution still have a bit lower employment rate than women at the bottom of the earnings distribution. The gender difference in employment is the lowest at the middle. During the 1997–2000 period, earnings were positively correlated with employment for both men and women, and the correlation was stronger for men. The positive relationship between these two variables became even stronger for both men and women in the 2000–2004 period, and the employment rate and earnings gradient of women was steeper than that of men. If earnings are positively correlated with human capital, then the above evidence suggests that human capital played a more and more important role in determining a person's probability of being employed. Because, on average, women have less human capital than men, strengthening the relationship between employment and human capital reduces the gender gap in both

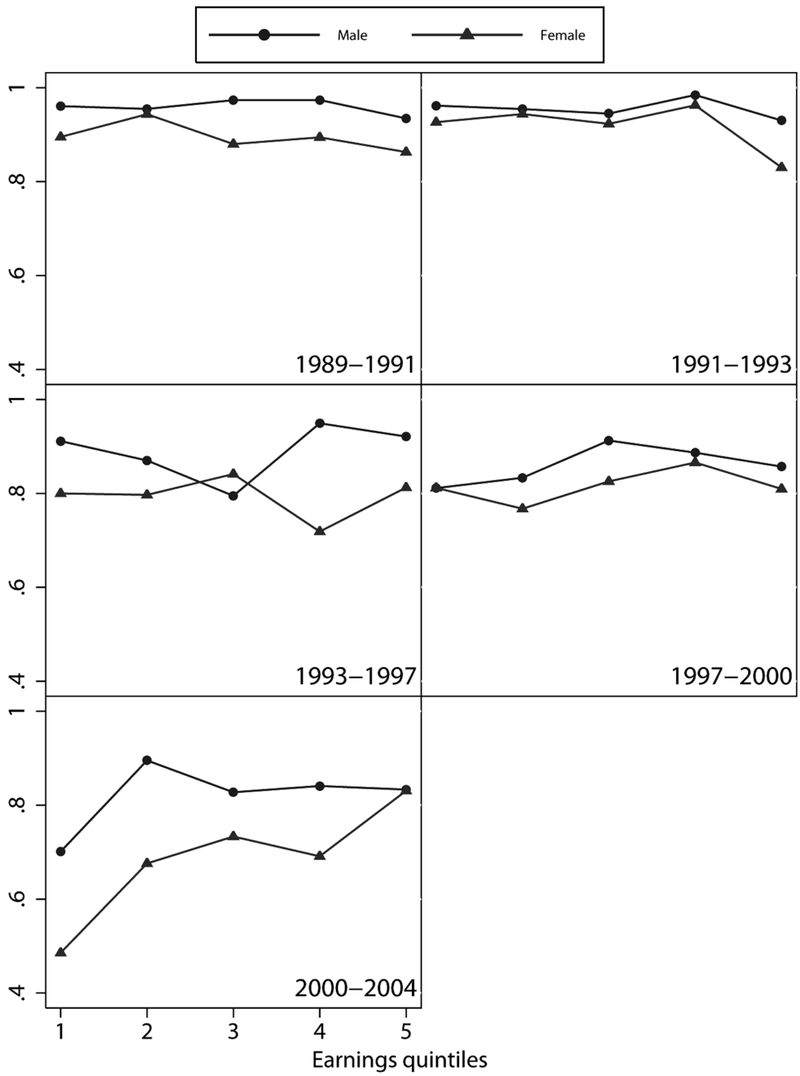

Figure 4. Employment rate of those who worked in the first wave, by earnings quantiles

observed and unobserved human capital. Therefore, the negative gap effect documented in table 5 does not necessarily suggest that the degree of labor market discrimination against women has decreased.

Although the graphic presentation is informative, it cannot reveal whether the higher job separation rate of lower-paid workers is due to their lower level of human capital or other factors that are positively correlated with their wages. For example, workers in SOUs do not only earn less than workers in the private sector but may also face a higher probability of being laid off due to the large scale of downsizing in the late 1990s. This generates a positive relationship between monthly earnings and employment rate. Another potential factor that can generate a similar relationship is the unbalanced economic growth across provinces. The large-scale SOE downsizing has a much stronger impact on workers in the northeastern provinces, such as Liao Ning and Hei Long Jiang, than on workers in the eastern coastal provinces, such as Jiangsu. Because workers in the northeastern provinces earn less than workers in the coastal provinces, without controlling for the province of residence, it is difficult to pin down the causes for the correlation between employment and earnings.

We use a probit model to address these issues. For workers who were employed in the first period of these 2-year panels, their probabilities of being employed in the second period are modeled as

$$\mathrm{Prob}(E_i = 1) = \Phi(Z_i\alpha), \tag{5}$$

where $E = 1$ if an individual is employed in the second period and 0 otherwise; $\Phi(\cdot)$ is the standard normal cumulative density function; and Z is a vector of control variables including age, education, relative height, registration type of work unit in the first survey, log of monthly earnings in the first survey, number of earners (except for the respondent) in the household in the second survey, and the province of residence. To capture the potential nonlinear relationship between employment and age, age is included as a series of dummy variables: between 16 and 29, 45 and 49, 50 and 54, and 55 and 60. These aged between 30 and 44 are used as the reference group. As a result, coefficients on age dummies should be interpreted as the difference between the corresponding age group and those aged between 30 and 44. Equation (5) is estimated for men and women separately.

Table 7 reports the estimation results.[9] Rather than reporting the coefficients of the probit estimation, we report the marginal effect. For continuous independent variables, it is the change in the probability for an infinitesimal change in the value of each independent variable. For dummy variables, it is the change in the probability for a discrete change in the value of each independent variable. For the sake of brevity, coefficients on the province of residence dummies are not reported.

[9] We have also tried to include marital status and the number of children under 12 years old as control variables. The coefficients on these two variables are not significant even at the 10% level in most of our regressions. The coefficients on other variables are not sensitive to the inclusion of these two variables either. For brevity, we only report the estimation results without these two control variables.

TABLE 7

THE DETERMINANTS OF EMPLOYMENT RATE

	Males					Females				
	1989–91	1991–93	1993–97	1997–2000	2000–4	1989–91	1991–93	1993–97	1997–2000	2000–2004
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Earnings in first wave	.0082	−.0017	.0794	.0641	.0441	.0120	−.0105	.0155	.0156	.0777
	(.0072)	(.0119)	(.0339)**	(.0315)**	(.0338)	(.0206)	(.0163)	(.0599)	(.0486)	(.0566)
Age ≥ 55	−.4420	−.3179	−.6527	−.4309	−.4063	−.6878	−.5838	−.7504	−.5201	−.4645
	(.1095)***	(.0987)***	(.0966)***	(.0849)***	(.0875)***	(.0983)***	(.1555)***	(.0941)***	(.1320)***	(.1056)***
50 ≤ age ≤ 54	−.0611	−.1617	−.0318	−.1882	−.1583	−.4211	−.3892	−.4223	−.2460	−.3957
	(.0561)	(.0832)*	(.0562)	(.0785)**	(.0724)**	(.0991)***	(.1177)***	(.1103)***	(.1005)**	(.0814)***
45 ≤ age ≤ 49	−.0450	−.0531	.0484	−.0698	−.0358	.0035	−.0252	−.1917	−.0657	−.2519
	(.0425)	(.0490)	(.0312)	(.0574)	(.0588)	(.0369)	(.0270)	(.0822)**	(.0573)	(.0832)***
Age ≤ 29	−.0079	−.0262	.0009	−.0321	−.1226	−.0217	.0127	.0053	.0056	−.0239
	(.0133)	(.0255)	(.0406)	(.0533)	(.0789)	(.0291)	(.0123)	(.0723)	(.0623)	(.1069)
Schooling	.0002	.0022	.0024	.0116	.0093	.0073	.0043	.0161	.0246	.0423
	(.0006)	(.0013)*	(.0042)	(.0049)**	(.0059)	(.0027)***	(.0021)**	(.0074)**	(.0065)***	(.0103)***
No. of earners	.0018	.0050	.0093	.0271	.0690	−.0101	.0070	.0373	−.0333	.0516
	(.0049)	(.0078)	(.0285)	(.0282)	(.0350)**	(.0185)	(.0094)	(.0454)	(.0383)	(.0545)

Relative height	.0011	−.0074	−.0209	.0009	.0034	−.0066	.0073	−.0083	−.0033	−.0031
	(.0029)	(.0053)	(.0159)	(.0171)	(.0189)	(.0119)	(.0057)	(.0302)	(.0238)	(.0233)
SOU employees	.0355	.0180	.1168	.2136	.0722	.1805	.0060	.0738	.0541	.1403
	(.0421)	(.0298)	(.0843)	(.0601)***	(.0568)	(.0614)***	(.0252)	(.1305)	(.0674)	(.0805)*
Collectives	.0043	.0153	.0505	.1410	−.0458	.1270	.0034	.0274	−.0051	.1601
	(.0103)	(.0140)	(.0491)	(.0313)***	(.0606)	(.0329)***	(.0210)	(.1189)	(.0664)	(.0678)**
Number of observations	668	513	355	465	428	581	497	320	389	346

Note. The sample consists of individuals aged between 16 and 60 who were in two adjacent surveys and worked in the first of the two surveys. The employment status (= 1 if employed) in the second survey is used as the dependent variable. Log earnings and the type of work unit registration refer to the job in the first survey. The residence of province is also used as a control variable. The values are the marginal effect of the corresponding variables. Numbers in parentheses are standard errors. The standard errors are corrected for the potential correlation among individuals of the sample household.

* Significant at the 10% level.

** Significant at the 5% level.

*** Significant at the 1% level.

Earnings in the first year have no significant impact on the employment status in the subsequent year except for men during 1993–97 and 1997–2000. Nevertheless, the sign of the coefficient is negative in 1991–93 and positive after that, which is consistent with the patterns shown in figure 4. Because we have controlled for education and age, the coefficient on first year earnings primarily reflects the impact of unobserved skills. Hence, the estimation results suggest that the patterns presented in figure 4 are mostly driven by observed skills.

Age is the most important determinant of employment, particularly for women. Its impact peaked in 1993–97, the period when *xia gang* was first implemented. After that, its impact gradually declined. In 1993–97, the employment rate of 55–60-year-old men was 65.3 percentage points lower than that of prime-age men, while the difference between 55–60-year-old women and prime-age women was 75.0 percentage points. The difference in employment rate between 50–54-year-olds and the prime-age workers was 3.2 percentage points and not significant for men, but was 42.2 percentage points for women and significant at the 1% level. These estimates suggest that it was the workers closer to retirement age whose jobs were the first to be destroyed. To further address the reason for job separation, figure 5 plots the labor market status of all individuals who worked in the first year of these 2-year panels. The figure shows that the proportion of jobless people increased considerably in 1997, particularly for women. Among those jobless women, only a small fraction was still looking for jobs; the majority of them were either retired or simply stayed at home. The evidence suggests that most nonemployed women left the labor market after separating from their previous jobs. Unfortunately, we do not have enough information to examine whether they withdrew voluntarily.

The impact of age on employment declined slightly after 1997. Nevertheless, it was still one of the most important factors, and it always had a stronger negative effect on women's employment. One potential reason for the gender difference in employment is that the mandatory retirement is 60 for men and 55 for women. This difference causes not only the lower employment rate of the 55–60-year-old women but also the lower employment rate of the 50–55 and even the 45–50-year-olds. This is because the difference in mandatory retirement age also leads to a difference in early retirement age. When nonperforming firms are forced to reduce their workforce, before laying off workers, they tend to encourage workers close to the retirement age to take earlier retirement first. As a result, women could be eligible for retirement as early as 45 at some SOUs. Hence, the higher job separation rate of older women does not necessarily imply that they are discriminated against in the

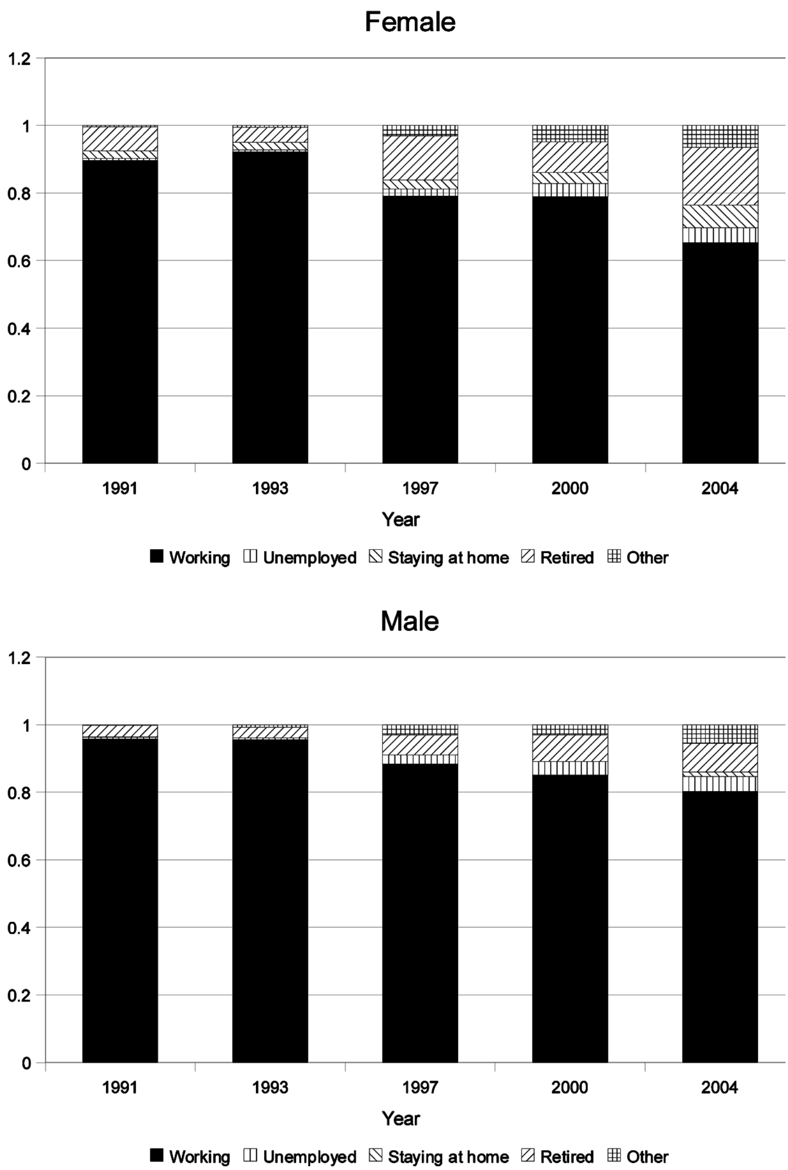

Figure 5. Labor market status of those who worked in the first wave

labor market. However, it does show that understanding the changes in gender difference in employment is crucial for us to understand the evolution of the gender earnings gap.

Years of education is another crucial factor for an individual's employment status. Again, its impact was larger for women. For example, a 1-year increase in schooling raised women's employment rate by 0.7 percentage points (significant at the 1% level) in the period 1989–91; it had no significant impact on men's employment during that period. Its impact increased to 4.2 (significant at the 1% level) percentage points in 2000–2004 for women, but its impact on men's employment was still not significant even at the 10% level during the same period. The rising impact of education on women's employment suggests that it became harder and harder for unskilled women to keep their jobs as economic reform progressed. If the impact of unobserved skills on employment had followed the same path, then the average level of women's unobserved skills should have increased relative to men's. As a result, the gender gap in unobserved skills should have narrowed over time. This is what was found from the analyses.

One potential reason for the higher job separation rate of unskilled women is the income effect. As has been documented in table 2, the average real monthly earnings for both men and women have more than quadrupled over the sample period. The fast earnings growth makes it possible for households to live on the incomes of one earner. As a result, some workers might voluntarily withdraw from the labor market. Due to their lower opportunity costs, the incentives to withdraw from the labor market are stronger for low-skilled workers. To examine this argument, we include the number of earners (excluding the respondent) as a control variable. Presumably, a person should have a much stronger incentive to work if he is the sole earner of the household. The coefficients on the number of earners are mostly insignificant. We have also tried to use the ownership of durable goods, such as a telephone and color TV, as a measure of household wealth; none of them support the argument that an individual's employment status is negatively correlated with household's wealth. Therefore, the higher exit rate of low-skilled women is unlikely to be due to the income effect.

Another potential reason for the higher job separation rate among low-skilled workers is the difference in education level across age groups. Because the average education level of older workers is lower than that of younger workers, the positive correlation between age and job separation rate might cause a spurious positive relationship between education and employment. If the difference in education level between older and younger female workers is larger than that between older and younger male workers, then the spurious relationship would be stronger for females. To test this hypothesis, we rerun the regression based on a sample of workers who are younger than 45 in the second survey. Presumably, age should have a limited effect on a person's employment for these workers, and the difference in education level across age groups should be smaller as well. Our estimation results show that education has a significant and positive effect on women's job retention rate in the 1997–2000 and 2000–2004 periods, but its impact for males never significantly differs from zero.[10] This evidence suggests that even among relatively young workers, education level still has a stronger positive impact on women's employment status.

V. Conclusion

In this study, we use a Chinese longitudinal data set, the China Health and Nutrition Survey, to examine the evolution of the gender earnings gap between 1989 and 2004. The longer period covered by the data set enables us to analyze whether the unprecedented large-scale downsizing of state-owned enterprises has worsened women's relative position in the labor market. Because economic reforms tend to have a larger negative impact on the employment of low-skilled workers and the average skill level of women is lower than that of men, the impact of economic reforms on women's relative status in the labor market can only be fully understood by examining earnings and employment simultaneously.

The results show that the gender earnings gap widened during the early urban reform period (1989–97) but that it has stabilized thereafter. The decomposition analysis shows that the stability was mainly achieved by a rise in the relative education level of employed women. Without these changes in observed characteristics, the gap would have increased by about 1.5 percentage points. Moreover, the examination on the determinants of employment status shows that the increase in the relative education level is the result of a higher exit rate from employment of less educated women. Unfortunately, the data do not contain enough information to allow an examination of whether these women withdrew from the labor market voluntarily.

Overall, the analyses show that women have shouldered more of the burdens of economic reforms in urban areas. The gender inequality in the labor market was mainly in the form of unequal pay in the early period of urban reform and in the form of unequal employment probability in the later period. Obviously, much more work needs to be done to explain why recent urban reforms had a stronger adverse impact on women's than on men's employment. How-

[10] For the sake of brevity, the full estimation results are not reported.

ever, the results of this study indeed show that to fully understand the impact of economic transition on gender inequality in the labor market, researchers need to examine earnings and employment simultaneously.

Appendix

TABLE A1
QUANTILE REGRESSION RESULTS, 10TH PERCENTILE

	1989 (1)	1991 (2)	1993 (3)	1997 (4)	2000 (5)	2004 (6)
Gender	.0632	.1024	.0864	.1048	.1386	.1767
	(.0402)	(.0266)***	(.0357)**	(.0463)**	(.0565)**	(.0388)***
Years of education	.0425	.0415	.0427	.0334	.0497	.0783
	(.0081)***	(.0046)***	(.0063)***	(.0082)***	(.0096)***	(.0075)***
Years of experience	.0325	.0248	.0309	.0262	.0221	.0188
	(.0095)***	(.0041)***	(.0059)***	(.0071)***	(.0096)**	(.0071)***
(Years of experience)^2	−.0439	−.0188	−.0416	−.0450	−.0255	−.0225
	(.0269)	(.0087)**	(.0121)***	(.0153)***	(.0205)	(.0157)
Relative height	−.0133	.0245	.0152	−.0032	.0498	.0472
	(.0238)	(.0171)	(.0209)	(.0274)	(.0341)	(.0204)**
SOU employees	.3195	.1600	−.1409	.0288	.1692	.0531
	(.1069)***	(.0850)*	(.0813)*	(.0818)	(.0897)*	(.0517)
Collectives	.2144	.0793	−.0220	−.1609	−.0077	−.1814
	(.1108)*	(.0865)	(.0798)	(.0860)*	(.0942)	(.0698)***
Controlled for residence of province	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	949	1,320	1,136	1,230	1,095	879

Note. The sample consists of individuals aged between 16 and 60. The log of real monthly average earnings is used as the dependent variable. The residence of province is also used as a control variable. Numbers in parentheses are standard errors. The standard errors are corrected for the potential correlation among individuals of the sample household.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

TABLE A2
QUANTILE REGRESSION RESULTS, 25TH PERCENTILE

	1989 (1)	1991 (2)	1993 (3)	1997 (4)	2000 (5)	2004 (6)
Gender	.0567	.0886	.1001	.1524	.1313	.1904
	(.0179)***	(.0181)***	(.0259)***	(.0365)***	(.0337)***	(.0571)***
Years of education	.0310	.0382	.0381	.0333	.0551	.0757
	(.0034)***	(.0031)***	(.0045)***	(.0065)***	(.0061)***	(.0112)***
Years of experience	.0280	.0254	.0252	.0246	.0247	.0125
	(.0044)***	(.0028)***	(.0041)***	(.0062)***	(.0062)***	(.0105)
(Years of experience)^2	−.0346	−.0221	−.0307	−.0329	−.0311	−.0145
	(.0125)***	(.0060)***	(.0090)***	(.0137)**	(.0134)**	(.0228)
Relative height	−.0027	.0204	.0059	.0120	.0342	.0393
	(.0108)	(.0114)*	(.0160)	(.0217)	(.0211)	(.0284)
SOU employees	.2777	.0103	−.0971	−.0129	.2300	.1150
	(.0514)***	(.0558)	(.0663)	(.0644)	(.0523)***	(.0734)
Collectives	.1804	−.0293	−.0290	−.1311	.0485	−.0179
	(.0519)***	(.0567)	(.0676)	(.0686)*	(.0560)	(.1005)
Controlled for residence of province	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	949	1,320	1,136	1,230	1,095	879

Note. The sample consists of individuals aged between 16 and 60. The log of real monthly average earnings is used as the dependent variable. The residence of province is also used as a control variable. Numbers in parentheses are standard errors. The standard errors are corrected for the potential correlation among individuals of the sample household.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

TABLE A3
MEDIAN REGRESSION RESULTS

	1989 (1)	1991 (2)	1993 (3)	1997 (4)	2000 (5)	2004 (6)
Gender	.0889	.0621	.1234	.1078	.1464	.0961
	(.0203)***	(.0133)***	(.0278)***	(.0254)***	(.0244)***	(.0330)***
Years of education	.0243	.0305	.0272	.0310	.0476	.0825
	(.0039)***	(.0023)***	(.0050)***	(.0047)***	(.0047)***	(.0066)***
Years of experience	.0166	.0162	.0184	.0207	.0216	.0255
	(.0049)***	(.0020)***	(.0042)***	(.0042)***	(.0042)***	(.0058)***
(Years of experience)^2	−.0080	−.0012	−.0188	−.0249	−.0241	−.0358
	(.0137)	(.0044)	(.0089)**	(.0094)***	(.0092)***	(.0127)***
Relative height	.0196	.0322	.0086	.0307	.0264	.0338
	(.0128)	(.0082)***	(.0169)	(.0150)**	(.0149)*	(.0171)**
SOU employees	−.0282	−.0893	−.2168	−.0101	.2137	.1412
	(.0594)	(.0420)**	(.0697)***	(.0437)	(.0374)***	(.0408)***
Collectives	−.1095	−.1062	−.1549	−.1526	−.0300	−.0585
	(.0604)*	(.0426)**	(.0714)**	(.0467)***	(.0402)	(.0570)
Controlled for residence of province	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	949	1,320	1,136	1,230	1,095	879

Note. The sample consists of individuals aged between 16 and 60. The log of real monthly average earnings is used as the dependent variable. The residence of province is also used as a control variable. Numbers in parentheses are standard errors. The standard errors are corrected for the potential correlation among individuals of the sample household.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

TABLE A4

QUANTILE REGRESSION RESULTS, 75TH PERCENTILE

	1989 (1)	1991 (2)	1993 (3)	1997 (4)	2000 (5)	2004 (6)
Gender	.0767	.1028	.1773	.1655	.1668	.1444
	(.0315)**	(.0205)***	(.0333)***	(.0241)***	(.0313)***	(.0290)***
Years of education	.0183	.0251	.0167	.0348	.0327	.0708
	(.0063)***	(.0037)***	(.0062)***	(.0046)***	(.0062)***	(.0060)***
Years of experience	.0103	.0133	.0230	.0206	.0089	.0247
	(.0072)	(.0030)***	(.0048)***	(.0039)***	(.0052)*	(.0048)***
(Years of experience)^2	.0093	.0012	−.0303	−.0233	−.0019	−.0304
	(.0199)	(.0065)	(.0100)***	(.0088)***	(.0114)	(.0107)***
Relative height	.0286	.0377	.0273	.0285	.0172	.0353
	(.0189)	(.0127)***	(.0200)	(.0140)**	(.0186)	(.0156)**
SOU employees	.0189	−.0978	−.5212	−.1704	.0487	.0932
	(.0952)	(.0651)	(.0844)***	(.0406)***	(.0465)	(.0351)***
Collectives	−.0143	−.0948	−.4950	−.3178	−.1248	−.1460
	(.0969)	(.0658)	(.0862)***	(.0445)***	(.0501)**	(.0499)***
Controlled for residence of province	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	949	1,320	1,136	1,230	1,095	879

Note. The sample consists of individuals aged between 16 and 60. The log of real monthly average earnings is used as the dependent variable. The residence of province is also used as a control variable. Numbers in parentheses are standard errors. The standard errors are corrected for the potential correlation among individuals of the sample household.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

TABLE A5

QUANTILE REGRESSION RESULTS, 90TH PERCENTILE

	1989 (1)	1991 (2)	1993 (3)	1997 (4)	2000 (5)	2004 (6)
Gender	.1324	.1569	.1628	.1512	.1548	.0688
	(.0508)***	(.0260)***	(.0430)***	(.0388)***	(.0480)***	(.0508)
Years of education	.0069	.0180	.0196	.0295	.0366	.0759
	(.0094)	(.0046)***	(.0074)***	(.0080)***	(.0103)***	(.0107)***
Years of experience	.0179	.0131	.0229	.0147	−.0009	.0209
	(.0105)*	(.0036)***	(.0061)***	(.0058)**	(.0079)	(.0086)**
(Years of experience)^2	−.0111	−.0031	−.0296	−.0132	.0178	−.0141
	(.0297)	(.0076)	(.0130)**	(.0127)	(.0181)	(.0190)
Relative height	.0378	.0111	.0498	.0119	.0004	.0802
	(.0279)	(.0154)	(.0244)**	(.0223)	(.0310)	(.0344)**
SOU employees	−.2326	−.3570	−.6787	−.3231	−.0187	.0243
	(.1310)*	(.0790)***	(.1046)***	(.0647)***	(.0701)	(.0633)
Collectives	−.1781	−.2517	−.5944	−.5200	−.0905	−.2304
	(.1369)	(.0803)***	(.1082)***	(.0718)***	(.0753)	(.0867)***
Controlled for residence of province	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	949	1,320	1,136	1,230	1,095	879

Note. The sample consists of individuals aged between 16 and 60. The log of real monthly average earnings is used as the dependent variable. The residence of province is also used as a control variable. Numbers in parentheses are standard errors. The standard errors are corrected for the potential correlation among individuals of the sample household.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

TABLE A6
EARNINGS REGRESSION RESULTS FOR MEN AND WOMEN

	1989 (1)	1991 (2)	1993 (3)	1997 (4)	2000 (5)	2004 (6)
A. Men:						
Schooling	.0174	.0243	.0178	.0227	.0353	.0653
	(.0062)***	(.0040)***	(.0061)***	(.0062)***	(.0085)***	(.0084)***
Experience	.0070	.0169	.0266	.0199	.0222	.0227
	(.0089)	(.0037)***	(.0055)***	(.0057)***	(.0082)***	(.0080)***
Experience^2/100	.0227	−.0070	−.0360	−.0284	−.0243	−.0288
	(.0249)	(.0076)	(.0115)***	(.0124)**	(.0172)	(.0170)*
Relative height	.0263	.0150	.0322	−.0001	.0376	.0602
	(.0212)	(.0146)	(.0212)	(.0198)	(.0269)	(.0254)**
SOU employees	.0280	−.1162	−.3374	−.2528	.1178	.0956
	(.1197)	(.0764)	(.0836)***	(.0610)***	(.0717)	(.0514)*
Collectives	−.0236	−.0746	−.2353	−.3850	−.0034	−.0707
	(.1218)	(.0781)	(.0869)***	(.0649)***	(.0776)	(.0739)
Constant	3.5005	3.8132	4.2835	4.4799	4.2166	4.3691
	(.1535)***	(.0945)***	(.1187)***	(.1036)***	(.1359)***	(.1364)***
Number of observations	476	696	605	663	603	495
B. Women:						
Schooling	.0229	.0364	.0371	.0473	.0468	.0955
	(.0072)***	(.0042)***	(.0069)***	(.0065)***	(.0096)***	(.0113)***
Experience	.0277	.0229	.0283	.0207	.0183	.0241
	(.0080)***	(.0036)***	(.0054)***	(.0056)***	(.0081)**	(.0093)***
Experience^2/100	−.0465	−.0187	−.0409	−.0229	−.0359	−.0330
	(.0223)**	(.0086)**	(.0119)***	(.0130)*	(.0189)*	(.0213)
Relative height	.0071	.0369	.0231	.0350	−.0142	.0317
	(.0222)	(.0145)**	(.0227)	(.0207)*	(.0305)	(.0246)
SOU employees	.0779	−.1008	−.1178	.0631	.2203	.0349
	(.0931)	(.0754)	(.0994)	(.0557)	(.0716)***	(.0693)
Collectives	.0061	−.1376	−.1155	−.1143	−.0356	−.2043
	(.0943)	(.0756)*	(.1000)	(.0600)*	(.0764)	(.0944)**
Constant	3.2815	3.5217	3.8236	3.8463	4.1585	3.9913
	(.1266)***	(.0883)***	(.1258)***	(.0968)***	(.1422)***	(.1659)***
Number of observations	473	624	531	567	492	384

Note. The sample consists of individuals aged between 16 and 60. The log of real monthly average earnings is used as the dependent variable. The residence of province is also used as a control variable. Numbers in parentheses are standard errors. The standard errors are corrected for the potential correlation among individuals of the sample household.
* Significant at the 10% level.
** Significant at the 5% level.
*** Significant at the 1% level.

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