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Changes in the structure of earnings during the Polish transition

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Abstract

We document changes in the structure of earnings during the economic transition in Poland. We find that inequality in labor earnings increased substantially from 1988 to 1996. A common view is that the reallocation of workers from a public sector with a compressed wage distribution, to a private sector with much higher wage inequality, accounts for the bulk of increased earnings inequality during transition (see, e.g., the models of Aghion and Commander (1999) [Aghion, Philippe, Commander, Simon, 1999. On the dynamics of inequality in the transition. *Economics of Transition* 7, 275–2898.] and Commander and Tolstopiatenko (1998) [Commander, Simon, Tolstopiatenko, Andrei, 1998. The role of unemployment and restructuring in the transition. In: Commander, Simon (Ed.), *Enterprise Restructuring and Unemployment in Models of Transition*. The World Bank, Washington, pp. 169–192.]). However, our decomposition of the sources of the increase in inequality suggests that this compositional effect accounts for only 39% of the increase. Fully 52% of the increase is due to the increase in the variance of wages within sectors. That is, earnings inequality within both the private and public sectors grew substantially, and by similar amounts. This is consistent with prior work suggesting that even state-owned enterprises in Poland moved towards competitive wage setting as they restructured (see, e.g., Pinto et al. (1993) [Pinto, Brian, Belka, Marek, Krajewski, Stefan, 1993. *Transforming state enterprises in Poland: evidence on adjustment by manufacturing firms*. *Brookings papers on Economic Activity*, 213–70.] Commander and Dhar (1998) [Commander, Simon, Dhar, Sumana, 1998. *Enterprises in the Polish transition*. In Simon Commander (Ed.), *Enterprise Restructuring and Unemployment in Models of Transition*. The World Bank, Washington, pp. 109–142.]).

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A substantial part of the increase in earnings inequality was between group, due largely to increased education premiums. However, changes in inequality within education–experience–gender groups account for about 60 percent of the increase in overall earnings inequality. The increases in within-group inequality were very different across skill groups, with much larger increases for highly educated workers. These patterns hold in both the private and public sectors, although increases in education premiums were somewhat greater in the private sector.

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1. Introduction

Poland experienced a dramatic change in its political and economic structures during the last decade. Its transition from a communist to a market economy began with a radical set of reforms in late 1989 and early 1990 known as the “big bang.” The communist government ended food price controls as it left power in August 1989, and the new Mazowiecki government implemented the Balcerowicz plan in January 1990, ending price controls on most other products. Other aspects of the reforms included reductions in state orders for manufactured goods and the imposition of hard budget constraints on state owned enterprises (SOEs). The hardening of budget constraints arose both through elimination of direct state subsidies and the reform of the National Bank in late 1991.¹

The privatization of SOEs in Poland began in 1990, but the pace of privatization has been slow compared to a number of other Eastern European countries. For instance, [Pinto et al. \(1993\)](#) report that, out of a sample of 75 of the 500 largest firms in Poland, only 3 had been privatized by June 1992 (two and one-half years into the transition).² According to [EBRD \(1995, 1997\)](#), the election of a left-of-center government in September 1993 slowed the pace further, but the new privatization law passed in July 1995 reversed this setback. Still, by the end of 1996, only about 44% of medium to large SOEs had begun the privatization process.

Nevertheless, the private sector’s contribution to GDP rose sharply during the transition (from 29% in 1989 to 60% in 1995) due largely to an explosion of small-scale entrepreneurship. In addition, there is strong evidence that hard budget constraints and import competition resulted in rapid adjustment by SOEs to the new market environment.

¹ [Commander and Dhar \(1998\)](#) report that total subsidies fell from 13% of GDP in 1989 to 2% in 1994. [IMF \(1994\)](#) reports that cash subsidies to SOEs declined from 4 percent of GDP in 1989 to below 1 percent of GDP by 1992.

² Twenty-four of the 75 SOEs had been “commercialized.” Essentially, this means that control was transferred from a workers’ council that could hire and fire managers, to a supervisory board that contains four members from the Ministry of Privatization and two members chosen by employees. But ownership remained with the Treasury. This was viewed as an intermediate step that would allow the firm to be restructured prior to privatization.

Pinto et al. (1993) provide a good discussion of the nature of these adjustments, which included massive labor shedding, changes in product lines and marketing strategies, and attempts to improve efficiency through investment.³

The transition also involved significant changes in labor market institutions. Constraints on layoffs and redundancies were significantly reduced. The unemployment rate rose from essentially zero in 1988 to a peak of 16.4% in 1994, and there has been massive inter-sectoral reallocation of labor.⁴ The rapid rise of the private sector—which is far less unionized than the public sector and much less subject to regulations in terms of wage setting—has also resulted in greater labor market “flexibility” in many dimensions. Very generous pensions led massive numbers of older workers to take early retirement during the early phase of the transition.

After a sharp contraction of output in 1990–1991, Poland experienced sustained economic growth, which became quite rapid in the mid-90s. As Keane and Prasad (2002) discuss, Poland was the greatest success story of the initial transition process. By 1999, its GDP was 22% above its pre-transition (1988) level, while even the best performing of the other transition countries had only recovered to within a few points (plus or minus) of their initial levels. At the same time, Poland experienced only a very modest increase in income inequality. Given its relative success, it is particularly important to document what happened during the transition process in Poland.

In this paper, we examine the evolution of the structure of labor earnings in Poland over the period 1985–1996 using micro data from the Polish Household Budget Surveys. The relatively long span of the dataset allows us to trace out changes for an extended period leading up to and following the “big bang.” We find that overall earnings inequality rose markedly during the transition period of 1989–1996. For instance, we estimate that the log 90–10 percentile differential for individual labor earnings increased from 0.97 in 1988 to 1.12 in 1996 (using a sample of individuals aged 18–60 for whom labor earnings is the primary income source). For men the increase was even greater, with the log 90–10 differential rising from 0.94 to 1.15.

We also conduct a detailed examination of the sources of the increase in earnings inequality. Prior to the transition, the wage structure in Poland was highly compacted, with wages of college-educated white-collar workers a little different from those of manual workers. A common view is that the rise of the private sector, in which there is competitive wage setting and, hence, a more unequal wage distribution, is the main source of increasing earnings inequality during transition. But our results contradict this view. In

³ Pinto et al. (1993) argue that managers of SOEs had two incentives to restructure: 1) so that they would be retained in the future (after privatization), and 2) the expectation that they would get cheap shares once privatization occurred. The notion that privatization per se would not lead to restructuring, but rather that the nature of managerial incentives and financial constraints are critical, is emphasized by Frydman and Rapaczynski (1994).

⁴ Coricelli et al. (1995) report that employment was 17.7 million in 1988. Of this, 9.6 million were in the public sector, 4.1 million were in private agriculture, 2.7 million were in worker cooperatives, and 1.3 million were in the nonagricultural private sector. By 1992, these figures were 6.6, 4.1, 1.1 and 3.8 million, respectively, along with 2.5 million unemployed. Thus, 2.5 million private sector jobs were created in just four years. Note that the labor force increased to 18.1 million, largely due to an inflow of women. Also, the share of the workforce in manufacturing fell from 37.2% in 1988 to 30.6% in 1996.

Poland, earnings inequality is indeed higher in the private sector (e.g., the log 90–10 earnings differential in 1996 was 1.19 in the private sector and 1.05 in the public sector), and the private sector share of (non-agricultural) employment did increase from 5% in 1988 to 39% in 1996. Still, we find that reallocation of labor from the public to the private sector accounted for only 39% of the total increase in earnings inequality (as measured by the change in the variance of log earnings).

The majority of the increase in earnings inequality during the Polish transition (52%) was due to increased variance of wages *within* both the public and private sectors. That is, earnings inequality within both the private and public sectors grew substantially, and by similar amounts.⁵ This is consistent with the view that even state-owned enterprises in Poland have engaged in substantial restructuring, as suggested by Pinto et al. (1993) and others. Consistent with our finding of increased earnings inequality within the public sector, Commander and Dhar (1998) report (p. 127) a substantial increase in the heterogeneity of wages across SOEs between 1990 and 1994, with those that performed better in terms of sales offering higher wages.⁶

We also find that educational wage premiums increased substantially. Nevertheless, the majority of the increase in overall earnings inequality (60%) in Poland is attributable to changes in within-group inequality. A striking result is that increases in within-group inequality were concentrated among workers with higher levels of formal education. This is quite different from the patterns documented for the U.S. and the U.K. of sharp increases over the last two decades in between-group inequality at all levels of education.⁷

In the large literature on rising wage inequality in the U.S. and U.K. in the 1980s, it has been common to attribute changes in wages between and within groups as being due to changes in the marginal product of observed and unobserved skills. But these interpretations rest on the assumption that wages closely reflect productivity, which may not be appropriate in the context of transition. Thus, for instance, increased wage premiums for college educated workers during the transition may reflect a process whereby the skills generated by education became more productive, due to changes in technology or technical efficiency (including better matching of workers to their most productive activities). Or it may reflect that more educated workers were more productive all along, so that their relative wages rose as wage rates came to be more closely aligned with marginal products. Or it could reflect some combination of both factors. Developing methods to sort out the relative importance of these mechanisms is an important topic for research. But we make no attempt to do that here, as we intend this paper to be purely

⁵ Exactly how similar is sensitive to the inequality measure chosen, but this in itself suggests it is not obvious that the increase in inequality was much greater in the private sector.

⁶ Since our data indicates sector of employment but does not identify a worker's specific firm, we cannot determine if the increased variability of wages in the public sector is within vs. between SOEs. This also prevents us from conducting an analysis like that of Brown and Earle (2004), who look at a sample of state owned and formerly state owned manufacturing firms in Russia and Ukraine, and find evidence of worker movement from less to more productive firms.

⁷ See, e.g., Juhn et al. (1993) and Gould, Moav and Weinberg (2001) for evidence from the U.S.; Machin (1996) and Machin and Van Reenen (1998) for the U.K.

descriptive. Nevertheless, we can ask whether the “stylized facts” uncovered by our descriptive analysis are consistent with existing models of transition. We do this in Section 5, where we conclude that existing models are not consistent with many of the stylized facts.

2. Review of prior research

Most earlier work on the Polish earnings distribution has relied on aggregate statistics that are released annually by the Polish Central Statistical Office (CSO). These aggregate statistics are described in detail in [Atkinson and Micklewright \(1992\)](#). Each September, starting in 1981, the CSO conducted a census of enterprises, and “the information requested of the enterprise was the total persons in a number of discrete earnings bands.” The CSO then published aggregate statistics such as total number of employees in various earnings bands, and deciles of the earnings distribution. From this data, it is possible to construct approximate measures on earnings inequality, such as approximate Gini coefficients. A number of other countries, such as Hungary, the former Czechoslovakia, and the former U.S.S.R., had similar data collection and reporting procedures for earnings.

Using such data, [Atkinson and Micklewright \(1992\)](#) compare the degree of earnings inequality across several communist countries in 1986. They obtain Gini values for Czechoslovakia (1987), Hungary, Poland and the U.S.S.R. of 0.197, 0.221, 0.242 and 0.276, respectively. And, based on the average earnings of individuals in the top and bottom deciles of the distribution, they report log 9–1 decile earnings differentials of 0.90, 0.97, 1.02 and 1.19, respectively. Thus, there was a clear ranking of inequality (consistent across both measures), with Czechoslovakia being the most equal, the U.S.S.R. the least equal, and Poland in the middle. As a point of comparison, these authors report a Gini coefficient and log 9–1 decile ratio of 0.267 and 1.17, respectively, for the U.K. in 1986. Thus, the earnings distribution in Poland prior to the transition was noticeably more compact than in the U.K. Based on the same data, [Atkinson and Micklewright \(1992\)](#) also calculate that earnings inequality in Poland declined over the period 1986–1989. They report Gini values of 0.242, 0.230, 0.212 and 0.207 for these years, and log decile differentials of 1.02, 1.02, 0.96 and 0.89.

[Rutkowski \(1996a\)](#) uses the same September earnings distribution survey to examine changes in the Polish earnings distribution during the transition. His calculations indicate that earnings inequality jumped dramatically in the early phase of the transition, with the Gini and log decile differential rising to 0.242 and 1.05 in 1991. By 1993, the last year of his study, these had risen further, to 0.257 and 1.11, respectively. Rutkowski also reports that earnings inequality was much greater in the private sector than the public sector in 1993, that the ratio of white collar to blue collar wages rose substantially in the transition, and that this ratio was much higher in the private sector. [Rutkowski \(1998\)](#) extends this analysis to 1995, by which time the Gini for earnings increased to 0.288 and the log decile differential to 1.22, a large increase in inequality over 1993. Recall that, for 1986, [Atkinson and Micklewright \(1992\)](#) reported a Gini of 0.242 and a log decile differential of 1.02. So the increase in (gross) earnings inequality for Poland from

1986–1995 implied by these data is a bit greater than what they report for the U.K. in the 1980s.⁸

Rutkowski (1996b) presents a cross-country analysis of changes in earnings inequality using similar data sources for several transition economies. These results indicate that, by 1993, the earnings distribution for Czechoslovakia had become very similar to that for Poland, while the Gini and log decile differential for Hungary had risen to 0.315 and 1.30 (more unequal than Poland). Thus, given the baseline figures from Atkinson and Micklewright (1992), it appears that both Czechoslovakia and Hungary experienced larger increases in earnings inequality from 1986 to 1993 than did Poland. Newell (2001) reports similar results.

As both Atkinson and Micklewright (1992) and Rutkowski (1996a,b) describe, there are a number of limitations of the September earnings survey data for Poland. First, the aggregate nature of the data may lead to approximation errors in inequality measures, and limits the type of analysis that can be performed. Second, the coverage of establishments is incomplete because small firms (i.e., less than 6 employees) are not sampled. This is especially a problem for the transition in Poland because, according to OECD (1998, p. 107), “Poland’s recent growth performance rests on a strong entrepreneurial basis, with many dynamic small and medium-sized enterprises (SMEs) and creations of new firms . . . SMEs make up the bulk of Poland’s 2.2 million registered non-agricultural enterprises. . . almost 90% [are] micro-enterprises (employing 1 to 5 persons).” Third, this data does not account for in-kind payments, which have been important in Poland. Fourth, this survey reports gross earnings. This creates comparability problems over time, because a progressive income tax (with rates up to 45%) was introduced in 1992. Failure to account for this will tend to exaggerate the measured increase in inequality of net earnings.

Another survey that has been used by some authors to examine recent changes in the Polish wage structure is the Polish Labor Force Survey that was introduced in 1992 (see, e.g., Newell and Socha, 1998). However, this survey is clearly not useful for understanding changes in wage inequality in the crucial early years of transition or changes relative to the pre-transition wage structure.

Existing work on other transition economies has focused mostly on the early years of transition (a recent exception is Newell, 2001). For instance, Orazem and Vodopivec (1995) analyze micro data from Slovenia and report that, from 1987 to 1991, wage inequality increased markedly, with returns to both education and experience rising over this period. Using grouped data, Flanagan (1995) finds that the returns to education rose while the returns to experience declined in the Czech Republic in the initial phase of transition. Brainerd (1998) reports that, from 1991 to 1994, the marginal return to a year of education almost doubled for workers in Russia. Garner and Terrell (1998) find that wage dispersion in the Czech and Slovak Republics increased in the early years of transition. Using data from the ILO, Freeman and Oostendorp (2000) have created the new Occupational Wages around the World file and, using this dataset, find that overall earnings inequality and skill differentials increased in transition economies during the 1980s and 1990s.

⁸ Note that, for both Poland and the U.K., the data are gross earnings.

3. The dataset

The CSO has been collecting detailed micro data on household income and consumption at least since 1978, using fairly sophisticated sampling techniques. In the HBS, the primary sampling unit is the household. A two-stage geographically stratified sampling scheme is used, where the first-stage sampling units are the area survey units and the second-stage units are individual households. The typical sample size is about 25,000 households per year. The CSO uses the data obtained from these household surveys to create aggregate tabulations that are then presented in their annual Statistical Bulletins, or *Surveys*.

The HBS contains detailed information on sources and amounts of income both for households and individuals within each household. Total income is broken down into four main categories: labor income (including wages, salaries and nonwage compensation); pensions; social benefits and other transfers; and other income. A key point is that the labor income data include measures of the value of in-kind payments from employers to workers, which have been an important part of workers' compensation in Poland and other transition economies. There were no taxes on personal income until 1992.⁹ After that year, we use net incomes in the analysis. The HBS also contains information on demographic characteristics of all household members and on labor earnings of all employed individuals in each household. Unfortunately, data on individual workers' earnings were not collected in 1993. Hence, the dataset we use in this paper in fact goes from 1985–1992 and then from 1994–1996.

The HBS includes a limited panel element—a part of the dataset contains households that are surveyed for four successive years before being rotated out of the sample. However, the attrition rate in the panel is significant and, in addition, the panel was changed completely a couple of times over the period 1985–1996. To maintain the representativeness of the sample and to use all of the information in the dataset, we treat the data as a repeated set of cross-sections.

The structure of the survey instrument and the sampling scheme were both kept essentially unchanged after the transition commenced. However, one major change was introduced in 1993 that has important implications for analyzing cross-sectional inequality. In order to improve survey response rates, in 1993 the CSO switched from quarterly to monthly data collection for the HBS. Since earnings are more variable at the monthly than the quarterly frequency, this change could have created a substantial increase in measures of cross-sectional earnings inequality. Indeed, as we show in the next section, failure to account for this change in survey frequency has quantitatively important effects on measures of earnings inequality.

In the Appendix, we develop a technique for adjusting the 1994–1996 earnings data for the increased variability that may be attributable to the shift from quarterly to monthly reporting. Our approach models earnings as the sum of a permanent or predictable component (determined by workers' education, age and other observable characteristics) and a mean zero idiosyncratic component. We then assume that the variance of the

⁹ Until 1992, some firms were levied an "excess wage tax," essentially a payroll tax imposed on part of a firm's total wage bill. The actual incidence of this tax is, of course, a complicated matter.

idiosyncratic component would not have jumped abruptly after the 4th quarter of 1992. Rather, we assume that the variance of idiosyncratic earnings varied smoothly over time (measured in months) according to a polynomial time trend. We estimate this polynomial trend, along with a dummy for post-1992 that captures the discrete jump in variance that occurred with the change to monthly earnings reporting. Then, at the individual level, we scale down the idiosyncratic component of the post-1992 earnings statistics to eliminate this jump in variance.

Our procedure for adjusting for the spurious increase in inequality stemming from the switch to the monthly reporting interval relies on access to the HBS micro data. In particular, the variance correction requires access to the data for an extended period of time. Our study is unique in that it is based on the HBS micro data for a long sample period extending from 4 years prior to the “big bang” to 7 years after. To our knowledge, no prior study of earnings inequality in Poland has adjusted for the change in survey design in 1993.¹⁰

We restrict our wage analysis sample to individuals between the ages of 18 and 60 who report that labor income is their principal income source. We deflate nominal wages using aggregate CPI data (1992Q4 = 100) for the survey quarter until the end of 1992 and for the survey month thereafter. Prior to 1993, there were 7 education categories reported for individuals in the survey. Beginning in 1993, two of these categories (basic vocational training and some high school) were combined into a single category; for consistency, we combine these categories in a similar manner for the 1985–1992 period. We also combined primary school and less than primary school into a single base category (among workers, the latter group is quite small), thereby yielding a total of 5 educational categories over the full sample.

The dataset contains sampling weights (at the household level) to correct for differences in survey non-response rates across household types and regions. We use these weights in our analysis, where appropriate, to maintain the cross-sectional representativeness of the sample. But none of our results differed much depending on whether or not we used the weights.

Table 1 reports sample means for some of the variables used extensively in our analysis.¹¹ The demographic characteristics of the cross-sectional samples remain

¹⁰ At the time we began our study, the Polish CSO had never before released the HBS micro data. Subsequently, the micro data for the first half of 1993 was released to the World Bank, and this data is used in Milanovic (1998). More recently, data for 1993–1996 have been obtained by researchers at the World Bank. A subsample of the HBS is now available through the Luxembourg Income Survey (LIS) for 1987, 1990 and 1992. Thus, no prior researchers have had access to the micro data for the entirety of the extended period that we examine. The HBS data is still being collected, but we decided to end our analysis in 1996 for two main reasons: First, our results suggest that inequality had reached a plateau in 1994–1996, so it appears that the main consequences of transition were worked out by that point. Second, it is very expensive to obtain data for additional years.

¹¹ In 1992, half the sample was used to test the new monthly survey; these data were considered unreliable and not made available to us. The sampling weights maintain the representativeness of that year’s data despite the fall in the sample size. Also note that the mean of the urban dummy rises sharply in 1992. This results from a reclassification of the “location” variable which we could not fully reconcile with that used in prior years. Hence, results for this variable should be interpreted with caution.

Table 1
Sample means for selected years: wage analysis sample

	1988	1989	1990	1991	1992	1994	1995	1996
College degree	0.09	0.09	0.09	0.10	0.11	0.12	0.11	0.12
Some college	0.01	0.01	0.01	0.00	0.00	0.03	0.03	0.03
High school	0.31	0.31	0.32	0.34	0.35	0.34	0.34	0.35
Vocational training	0.38	0.39	0.40	0.39	0.39	0.38	0.39	0.38
Primary school	0.21	0.21	0.18	0.16	0.14	0.13	0.13	0.12
Experience	20.95	21.39	21.26	20.87	21.25	20.44	20.66	20.55
Experience ²	5.37	5.54	5.47	5.24	5.36	5.10	5.19	5.17
Male	0.55	0.55	0.55	0.55	0.55	0.53	0.53	0.52
Urban	0.64	0.63	0.64	0.65	0.79	0.74	0.72	0.72
Private sector	0.05	0.06	0.09	0.13	0.18	0.31	0.35	0.39
Number of observations (households)								
1985	17,927	1988	21,114	1991	19,025	1995	26,316	
1986	19,666	1989	20,787	1992	7259	1996	26,165	
1987	21,574	1990	20,685	1994	26,095			

The wage analysis sample includes workers between the ages of 18 and 60 who report a positive wage and who report that labor income was their primary source of income.

quite stable during and after the transition. There is a steady increase in average levels of educational attainment in the 1990s, largely reflecting higher education levels of new cohorts entering the workforce. The distribution of employment among men and women is relatively stable, although there is a slight increase in the share of women in total employment after 1994. Private enterprises accounted for less than 10% of total employment before the transition but this proportion had grown to about 40% by 1996.

4. Earnings inequality

In this section, we examine the evolution of earnings inequality, using data for individual workers. For the years 1994–1996, we use earnings measures that are adjusted for the increase in idiosyncratic variance that occurred with the shift to a monthly reporting period (see the Appendix for details).

4.1. Measures of overall inequality

The first panel of Table 2 reports 90–10 and 75–25 percentile differentials of log earnings for all workers. Earnings inequality is quite stable in the pre-transition years 1985–1988, followed by a period of rapid growth in inequality that begins in 1989, the first year of transition. Between 1988 and 1989, the 90–10 differential increases sharply, going from 0.97 to 1.04. A further significant increase occurs from 1991 to 1992, followed by a moderate increase through 1996. The total increase in inequality from 1988 to 1996, as measured by the 90–10 differential, is about 15%, a sizeable increase over an 8-year

Table 2
Percentile differentials

	All workers					
			Unadjusted data		Wage residuals	
	90–10	75–25	90–10	75–25	90–10	75–25
1985	0.96	0.49	–	–	0.78	0.40
1986	0.97	0.50	–	–	0.78	0.40
1987	0.96	0.49	–	–	0.78	0.40
1988	0.97	0.50	–	–	0.82	0.42
1989	1.04	0.53	–	–	0.89	0.46
1990	1.03	0.52	–	–	0.90	0.46
1991	1.03	0.53	–	–	0.76	0.44
1992	1.08	0.56	–	–	0.92	0.46
1994	1.11	0.56	1.15	0.58	0.92	0.46
1995	1.12	0.58	1.16	0.60	0.92	0.46
1996	1.12	0.59	1.15	0.61	0.91	0.46

	Public sector		Private sector		Men		Women	
	90–10	75–25	90–10	75–25	90–10	75–25	90–10	75–25
	1985	0.95	0.49	1.03	0.53	0.92	0.45	0.84
1986	0.96	0.49	1.01	0.54	0.91	0.46	0.86	0.44
1987	0.96	0.49	1.08	0.54	0.91	0.45	0.85	0.43
1988	0.96	0.49	1.04	0.56	0.94	0.47	0.85	0.44
1989	1.03	0.53	1.14	0.61	1.00	0.51	0.97	0.51
1990	1.02	0.52	1.12	0.60	1.01	0.52	0.98	0.50
1991	1.00	0.51	1.16	0.61	1.02	0.53	0.95	0.48
1992	1.06	0.55	1.15	0.57	1.07	0.55	0.99	0.51
1994	1.06	0.54	1.19	0.61	1.14	0.58	0.99	0.50
1995	1.08	0.57	1.19	0.57	1.13	0.60	0.99	0.50
1996	1.05	0.55	1.19	0.61	1.15	0.59	1.00	0.51

The percentile differentials reported above are for log quarterly (1985–1992) or monthly (1994–1996) earnings. The data for 1994–1996 are adjusted for the change in survey frequency after 1992, unless indicated otherwise. Wage residuals are from annual OLS regressions of earnings on a constant, four education dummies, experience and its square, and dummies for gender, urban residence, and employment in the private sector.

period. The increase in the 75–25 percentile differential is also quite substantial, from 0.50 in 1988 to 0.59 in 1996.¹²

It is notable that the 90–10 differential seems to reach a plateau in 1994–1996. This suggests that, by extending our analysis to that point, we have largely captured the main inequality increasing effects of the transition. Even by 1996, however, the wage structure remains considerably more compressed than in the U.S. For instance, in 1991, the 90–10 differential for full-time workers in the U.S. was close to 1.75 (Gottschalk and Smeeding, 1997). Interestingly, however, wage differentials are greater in Poland than in certain continental European countries such as Germany in the 1990s (see Prasad, 2004).

¹² Finer breakdowns of the percentile differentials (not shown here) indicated that, during the transition, inequality above the median, as measured by the 90–50 and 75–50 differentials, increased slightly more than inequality below the median (the 50–10 and 50–25 differentials).

In [Table A1](#), we present an alternative measure of inequality—the Gini coefficient. It is evident that the patterns of changes in inequality revealed by the evolutions of the percentile differentials are, in general, quite similar to those indicated by the Gini coefficients. For instance, the Gini coefficient for the full sample rises sharply from 0.221 in 1988 to 0.259 in 1994 and then remains relatively flat through 1996.

Our results for earnings inequality in the HBS differ in a number of important ways from the results of [Atkinson and Micklewright \(1992\)](#) and [Rutkowski \(1998\)](#), who used aggregated data from the census of enterprises conducted by the CSO each September. We do not find evidence of the sharp drop in inequality from 1986–1989 that they report. And we find more modest increases in inequality after 1989. Starting from a base year of 1986 (when our figures roughly agree) and ending with 1995 (the final year in their analysis) we obtain a 16% increase in the log 90/10 differential and a 0.042 point increase in the Gini, while their figures imply 30% and 0.076 point increases, respectively. Given the much more representative population coverage of the HBS data, we view our results as more accurate.

As discussed earlier, we adjust the earnings data for 1994–1996 to account for the change in survey frequency. In the second panel of [Table 2](#), we show the percentile differentials for 1994–1996 with unadjusted data. Clearly, the adjustment makes a significant difference to the absolute level of inequality, although the profile of stable inequality over the period 1994–1996 is unaffected by whether the data are adjusted. However, adjusting for the change in survey frequency is clearly important to accurately measure the change in inequality over the full sample period.

We note that [Newell and Mieczyslaw \(1998\)](#), using the Labor Force Survey, find a comparable increase in earnings inequality from 1992–1996, with virtually all of this increase occurring between 1992 and 1994. This is very similar to what we find using our adjusted data for 1994–1996, suggesting that our adjustment procedure does not introduce any spurious inequality dynamics. But the dataset that we use has the distinct advantage that, unlike the Labor Force Survey which commenced in 1992, it includes data from earlier years; this is crucial since most of the increase in wage inequality seems to have occurred in the first few years of transition.

4.2. Changes in earnings inequality in the public vs. private sectors

In this section, we compare the earnings distributions in the state and private sectors. The first four columns in the lower panels of [Table 2](#) show percentile differentials separately for workers employed in the public and private sectors. In absolute terms, inequality is higher in the private sector than in the public sector. However, whether inequality increased more in the private or public sector during the transition is ambiguous. For instance, from 1988 to 1996, the 90–10 differential rises from 0.96 to 1.05 in the public sector (+9%) and from 1.04 to 1.19 in the private sector (+15%). However, the 75–25 differential actually increases slightly more in the public sector (+6%) than in the private sector (+5%).

It is also of interest to compare levels of earnings in the private vs. public sectors. If earnings differ substantially between the sectors, then the allocation of workers between them can alter aggregate measures of inequality. For instance, in the model of [Aghion and](#)

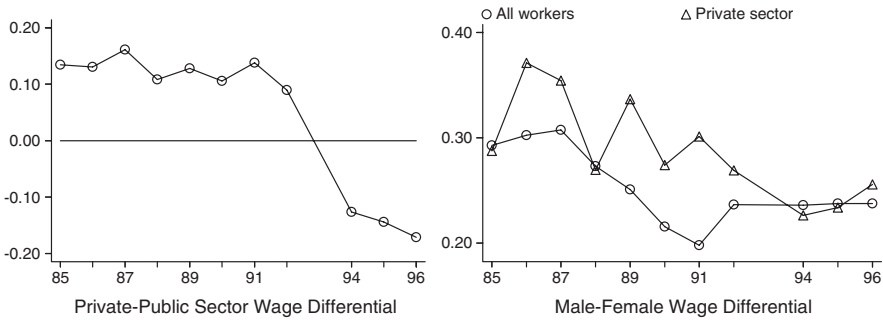


Fig. 1. Wage differentials.

Commander (1999), reallocation of workers from a low wage state sector to a high wage private sector is one factor driving up inequality. Fig. 1 (left panel) plots the differential between (unconditional) median private and public sector earnings. Surprisingly, the median earnings in the private sector dropped from about 10% above that in the public sector in 1992 to about 12% below in 1994.¹³ Compositional effects are important in driving these changes. Somewhat surprisingly, it is the low education workers who have primarily shifted into the emerging private sector.

Next, in order to better understand the changes in shape of the earnings distribution in each sector, we examine kernel density estimates. Fig. 2 presents kernel density estimates for (log) real earnings for 1988 and 1996 for each sector.¹⁴ To focus on changes in shape, we scaled earnings to have the same mean in each sector in each year.¹⁵ It is obvious from Fig. 2 that inequality increased in both the public and private sectors. In the private sector, the mass near the mode of the distribution clearly drops from 1988 to 1996. At same time, the distribution becomes skewed to the right. Crucially, the loss in mass near the mode is not obviously greater in the private sector than in the public sector. However, there is much more mass shifted into the extreme tails. This explains our earlier finding that when one looks at 75–25 differentials it appears that inequality increased slightly more in the public sector, but when one looks at 90–10 differentials it appears that inequality increased more in the private sector.

To sum up, the surprising finding of this section is that, by some measures, inequality grew as much in the state sector as in the private sector. This appears consistent with prior work by Pinto et al. (1993), Commander and Dhar (1998) and others, suggesting substantial restructuring by SOEs, such that wages in the public sector more closely reflect productivity.

¹³ Earnings data from the LFS reveal a similar pattern. Newell and Mieczyslaw (1998) report a mean private sector wage premium of 5 percent in 1992, falling to a negative differential of 10 percent by 1996.

¹⁴ An Epanechnikov kernel with a bandwidth of 0.05 was used for the kernel density estimation. We also computed optimal bandwidths—these were generally in the range of 0.04–0.06 and made little difference to the density plots.

¹⁵ Sensible inequality measures like percentile ratios, Gini coefficients and variances of log earnings are invariant to proportional scaling of earnings, since a change in the denomination of currency units should not alter inequality.

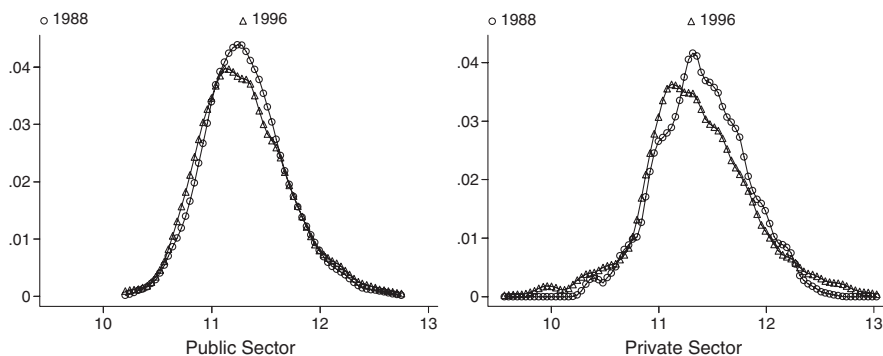


Fig. 2. Kernel density estimates, 1988 vs. 1996.

4.3. Effects of changes in the structure of employment on earnings inequality

In this section, we decompose the increase in overall earnings inequality into components attributable to changes in the composition of employment across sectors or industries versus increases in inequality within sectors or industries. Consider the following decomposition:

$$\sigma_t^2 = \sum_j s_{jt} \sigma_{jt}^2 + \sum_j s_{jt} (w_{jt} - \bar{w}_t)^2 \quad (1)$$

where σ_t^2 is the cross-sectional variance of log hourly earnings, s_{jt} is the employment share of sector j , σ_{jt}^2 is the within-sector variance of earnings, w_{jt} is sector j mean earnings, \bar{w} is grand mean earnings in the sample, and the subscript t is a time index. Using this formula, the change in variance over time can be decomposed into changes attributable to within- and between-sector components, as well as composition effects within and between sectors.

The top panel of Table 3 shows this variance decomposition based on the state and private sectors. The overall increase in log earnings variance from 1988 to 1996 is 7.00 (the figures in the table are multiplied by 100). Of this, 3.63 points, or 52%, is due to increases in variance *within* the state and private sectors. 2.71 points or 39% of the increase in variance is attributable to the shift of workers from the (relatively low variance) public sector to the private sector. Note that the two variance components that arise due to the differences in mean earnings between the two sectors (last two columns) are of minor importance.

These results suggest, somewhat surprisingly, that the shifting of workers from the state to the private sector, while important, is not the main factor driving increased earnings inequality during the Polish transition. Rather, *within* sector increases in earnings inequality constitute the most important factor driving the increase in overall inequality.¹⁶

¹⁶ Of course, there are two ways to decompose the change in variance, depending on whether one calculates the within sector component of the change in variance (i.e., the sum of the changes in sector specific variances weighted by sectoral employment shares) using the base period or terminal period employment shares. We use base period+shares as weights, which means the change in variance in the state sector dominates this term. If we use the terminal period shares instead, our conclusion that this is the most important term in the decomposition is strengthened.

Table 3
Effects of sectoral shifts on changes in wage inequality (Variance decomposition)

Period	Total change in variance	Within sector		Between sector	
		Change in variance	Composition effect	Change in variance	Composition effect
1988–1996	7.00	3.63	2.71	0.44	0.23
1988–1992	2.69	2.30	0.32	0.00	0.06
1992–1994	3.61	1.59	1.73	0.17	0.12
1994–1996	0.70	–0.26	0.66	0.27	0.05

Period	Total change in variance	Within industry		Between industry	
		Change in variance	Composition effect	Change in variance	Composition effect
1988–1996	7.00	6.76	–0.41	0.89	–0.23
1988–1992	2.69	2.39	0.21	0.21	–0.10
1992–1994	3.61	3.52	–0.21	0.19	0.09
1994–1996	0.70	0.85	–0.41	0.49	–0.22

Variance (of log wages) and its components multiplied by 100. For the results in the first panel, workers were classified into one of two sectors—private or public. For the results in the second panel, workers were classified into 13 broadly defined sectors. These sectors and their employment shares in 1988 and 1996 are as follows: manufacturing and mining (37.2, 30.6); construction (8.7, 8.0); agriculture, forestry, and fishing (11.7, 4.7); transport and communications (8.5, 8.5); trade, hotels, and restaurants (9.6, 11.5); real estate and housing (3.7, 3.0); science and technology (1.0, 0.7); education and culture (8.6, 9.7); health and hygiene (5.6, 8.5); recreation and tourism (0.7, 1.2); public administration (2.5, 7.2); finance and insurance (0.9, 2.5); and other services (1.2, 3.8).

A key factor driving this result is that earnings inequality in the state sector has grown substantially, as we documented in Sections 4.1 and 4.2. In 1996, the state sector still accounted for almost two-thirds of nonagricultural employment in Poland, so developments there remain critical for the shape of the overall earnings distribution. Despite the slow pace of privatization in Poland, work by [Pinto et al. \(1993\)](#), [Commander and Dhar \(1998\)](#) and others suggests that substantial restructuring of SOEs has nevertheless occurred, as managers responded to removal of government subsidies and import competition, thus leading the state sector closer to competitive wage setting.

Table 3 also breaks down the increase in overall earnings variance into sub-periods. Note that there were sharp increases in earnings variance from 1988 to 1992 and from 1992 to 1994, with a subsequent much more moderate increase from 1994 to 1996. It is interesting that composition effects were of almost no importance in the early transition period (1988–1992). But in the most recent period, the shifting of workers from the public to the private sector was the main factor driving increased earnings inequality. As shown in [Table 2](#), in 1994–1996 earnings inequality reached a plateau in both the public and private sectors. Indeed, within sector earnings dispersion appears to have fallen marginally from 1994 to 1996. Thus, compositional effects dominate the (modest) growth in earnings inequality during this latter period.¹⁷

¹⁷ The private sector share of total employment rose by 13 percentage points from 1992 to 1994 and by an additional 8 percentage points from 1994 to 1996 (see [Table 1](#)).

The composition of employment by industry also changed dramatically from 1988 to 1996. For instance, as noted in the footnote to Table 3, the share of manufacturing dropped from 37.2% to 30.6%. In the bottom panel of Table 3, we present a variance decomposition for 13 broadly-defined industrial sectors of the economy. The key result is that virtually all of the increase in overall earnings variance is attributable to within-industry increases in variance. The between-industry component of the change in variance (due to changes in the relative means of earnings across industries) is positive in all three sub-periods, but is roughly offset by within- and between-industry composition effects. Thus, industry employment shifts do not seem to have played much of a role in influencing patterns of overall earnings dispersion.¹⁸

In summary, growth of earnings inequality within sectors and industries is the main source of increased overall earnings inequality. Labor reallocation from the public sector to the private sector has also contributed importantly to the rise in overall inequality. However, despite the large shifts in industry employment shares, inter-industry labor flows do not appear to have contributed directly to the rise in earnings inequality. These two sets of results are reconciled by the fact that, while industry shares of total employment changed significantly for only a few industries, increases in the shares of private sector employment *within* each industry were very large.¹⁹

4.4. Within-group earnings inequality

In this section, we examine changes in within-group inequality. In Table 4, we report percentile differentials for workers in each of four educational groups. Prior to the transition, both the 90–10 and 75–25 percentile differentials were not too dissimilar across these groups. However, workers with college degrees experience by far the greatest increase in inequality during the transition, with the 90–10 differential rising by 0.18 and the 75–25 differential rising by 0.12 from 1988 to 1996. For workers with high school degrees, the corresponding increases are 0.12 and 0.08, respectively. For workers with only basic vocational training or a primary school degree, the increases are much more modest. For instance, changes in the 90–10 differential for workers with vocational training (0.05) or a primary school degree (0.09) are less than half of the corresponding change for workers with a college degree. Based on the 75–25 differential, the differences across educational groups in inequality growth are even greater. Thus, increases in within-group inequality seem to be a prominent feature of the transition mainly for highly educated workers.

One might surmise the explanation of this result is that better-educated workers were more likely to migrate to the private sector, where inequality is higher. In fact, private

¹⁸ We also recomputed this decomposition restricting the sample to workers in the private sector and found that, again, virtually all of the increase in log wage variance could be attributed to changes in within-industry inequality rather than composition effects. Thus, within industry wage variation appears to dominate overall wage variation and both appear to have evolved in a similar pattern.

¹⁹ For instance, the share of manufacturing and mining in total employment fell from 37.2% in 1988 to 30.6% in 1996. Over this period, the share of private sector employment in this industry rose sharply, from 5.9% to 43.5%. Similarly, while the fraction of workers in the trade sector rose from 9.6% in 1988 to 11.5% in 1996, the share of private sector employment within this industry jumped from 4.2% to 70.4% over this period.

Table 4
Within-group inequality: percentile differentials

	College degree		High school	
	90–10	75–25	90–10	75–25
1985	0.87	0.46	0.89	0.46
1986	0.88	0.44	0.90	0.46
1987	0.87	0.46	0.92	0.47
1988	0.96	0.50	0.91	0.46
1989	0.95	0.50	0.98	0.51
1990	0.95	0.48	0.95	0.49
1991	0.98	0.50	0.95	0.49
1992	1.03	0.54	1.00	0.52
1994	1.14	0.62	1.03	0.53
1995	1.15	0.62	1.04	0.54
1996	1.14	0.62	1.03	0.54
	Vocational training		Primary school	
	90–10	75–25	90–10	75–25
1985	0.99	0.50	0.90	0.47
1986	0.99	0.50	0.94	0.49
1987	0.98	0.50	0.93	0.47
1988	0.98	0.50	0.93	0.47
1989	1.04	0.52	0.97	0.51
1990	1.03	0.52	0.93	0.48
1991	1.00	0.51	0.91	0.45
1992	1.05	0.55	0.93	0.49
1994	1.05	0.55	0.92	0.47
1995	1.04	0.54	0.99	0.48
1996	1.03	0.55	1.02	0.50

The percentile differentials reported above are for log quarterly (1985–1992) or monthly (1994–1996) earnings. The data for 1994–1996 are adjusted for the change in survey frequency after 1992.

sector employment shares of all education groups rose sharply during the transition. But, somewhat surprisingly, movement into the private sector was most pronounced for workers with *lower* levels of education.²⁰ Hence, differential patterns of reallocation of labor across the public and private sectors cannot explain our finding. Indeed, we also found that increases in within-group inequality were greater for better-educated workers in each sector separately.

Next, we examine the evolution of inequality within broadly defined (synthetic) experience groups. Table 5 (top panel) reports percentile differentials for groups of workers with different experience levels. There are fairly significant increases in inequality for all groups, consistent with the plausible interpretation of these increases as reflecting time effects. For instance, from 1988 to 1996, the 90–10 differential rises by about 0.2 for all experience levels. Over the same period, the 75–25 differential rises by about 0.1 for all experience groups.

²⁰ The percent of workers in each education category employed in the private sector in 1988 and 1996 are as follows: college degree (2.9 in 1988, 19.4 in 1996); some college (5.3, 20.3); high school (3.4, 33.9); vocational training (6.6, 48.3); and primary school (3.8, 44.7).

Table 5
Wage inequality within synthetic experience groups and cohorts

Experience group (years):	90–10 differential					75–25 differential						
	0–10	11–20	21–30	31–40	41–50	0–10	11–20	21–30	31–40	41–50		
1985	0.93	0.92	0.93	0.91	0.89	0.46	0.47	0.46	0.48	0.42		
1986	0.99	0.95	0.93	0.94	0.84	0.49	0.49	0.47	0.47	0.42		
1987	0.98	0.97	0.92	0.89	0.82	0.50	0.49	0.46	0.45	0.43		
1988	0.93	0.96	0.94	0.92	0.81	0.44	0.50	0.48	0.47	0.45		
1989	1.03	1.03	1.01	0.99	0.90	0.52	0.53	0.51	0.51	0.49		
1990	1.05	1.03	1.02	0.98	0.91	0.53	0.53	0.52	0.49	0.50		
1991	0.99	1.04	1.01	0.99	0.99	0.49	0.53	0.52	0.50	0.49		
1992	1.05	1.09	1.08	0.98	1.03	0.55	0.56	0.54	0.54	0.55		
1994	1.07	1.12	1.14	1.06	0.97	0.52	0.56	0.58	0.54	0.54		
1995	1.04	1.14	1.15	1.11	1.00	0.49	0.59	0.59	0.58	0.54		
1996	1.05	1.10	1.15	1.11	1.03	0.54	0.59	0.60	0.59	0.53		
Cohort:	1926– 1935	1936– 1945	1946– 1955	1956– 1965	1966– 1975	>1976	1926– 1935	1936– 1945	1946– 1955	1956– 1965	1966– 1975	>1976
1985	0.92	0.90	0.92	0.92	0.79	–	0.48	0.45	0.47	0.47	0.37	–
1986	0.93	0.92	0.94	0.97	0.77	–	0.46	0.47	0.48	0.50	0.39	–
1987	0.86	0.91	0.94	0.96	0.80	–	0.45	0.45	0.48	0.49	0.40	–
1988	0.89	0.96	0.94	0.96	0.75	–	0.47	0.47	0.48	0.48	0.40	–
1989	0.97	1.02	1.01	1.03	0.89	–	0.51	0.52	0.52	0.51	0.45	–
1990	0.98	1.01	1.02	1.02	0.96	–	0.51	0.51	0.52	0.52	0.46	–
1991	1.00	1.03	1.01	0.99	0.93	–	0.56	0.52	0.52	0.51	0.45	–
1992	1.06	1.02	1.09	1.05	0.99	–	0.54	0.55	0.54	0.55	0.49	–
1994	–	1.07	1.11	1.13	1.05	0.78	–	0.53	0.55	0.57	0.51	0.28
1995	–	1.11	1.11	1.16	1.06	0.79	–	0.57	0.58	0.59	0.51	0.33
1996	–	1.12	1.11	1.12	1.07	0.88	–	0.58	0.57	0.60	0.56	0.37
					1985– 1988	1988– 1992	1992– 1996	1985– 1988	1988– 1992	1992– 1996		
Avg. change within experience groups:					0.00	0.13	0.04	0.01	0.08	0.02		
Avg. change within cohorts:					0.02	0.14	0.06	0.01	0.08	0.04		

The percentile differentials reported above are for log quarterly (1985–1992) or monthly (1994–1996) earnings. The data for 1994–1996 are adjusted for the change in survey frequency that occurred after 1992. Cohorts are defined on the basis of year of birth. Results are reported only for cell sizes greater than 100.

We also examined the evolution of overall inequality within birth cohorts. The middle panel of Table 5 reports log percentile ratios for synthetic cohorts defined on the basis of birth year. It is, of course, impossible to separate out age effects from time effects by looking at these differentials for any one cohort.²¹ But the fact that all cohorts experienced larger increases in inequality during the transition than during the pre-transition period, with a substantial fraction of this increase occurring between 1988–1989 and 1992–1994, suggests that time effects are important.

²¹ Typically, inequality tends to rise over the life cycle within a given cohort as employment histories, cumulative effects of individual productivity shocks, and other factors drive up within-cohort wage dispersion.

The cohort of younger workers, born between 1966 and 1975, experiences an increase of the 90/10 ratio from 0.75 in 1988 (when its oldest members are 22) to 0.99 in 1992 (when its oldest members are 26). Much of this increase is presumably due to age effects. But it is worth noting that private sector employment is much greater for this cohort than for the older ones.²²

While it is impossible to (nonparametrically) disentangle age, cohort and time effects on overall inequality, a plausible way to identify time effects was suggested by Juhn et al. (1993, p. 424–426). If we take the average (across age/experience groups) of changes in inequality between any two points in calendar time, we get changes that combine cohort and time effects. On the other hand, if we take the average (across cohorts) of changes in inequality between any two points in calendar time we get changes that combine age and time effects. So these two measures of changing inequality have time effects in common. If they move together, it is plausible that they do so because time effects are the dominant factor (rather than because age and cohort effects just happen to be equal). The bottom panel of Table 5 shows that changes in inequality (for different sub-periods) are indeed quite similar when they are averaged across cohorts vs. across age/experience groups, suggesting that much of the increase in the dispersion of the overall earnings distribution can plausibly be attributed to time effects.²³

4.5. Residual earnings inequality

Another approach to examine within-group wage inequality is to regress earnings on observed attributes such as gender, education, and experience and to examine the dispersion of the wage residuals. These earnings residuals arguably control for between-group differences across many different group characteristics and indicate the evolution of inequality within narrowly defined groups.

We report percentile differentials for log earnings residuals in the last (top right) panel of Table 2.²⁴ Comparing these to the ratios for log earnings themselves, we see that about four-fifths of overall earnings inequality is within-group. Changes in within-group inequality also account for a substantial fraction of the increase in overall inequality. For instance, the 90–10 percentile differential for earnings residuals goes from 0.82 in 1988 to 0.91 in 1996, which accounts for about 60% of the increase in total earnings inequality over that period.

²² The share of total employment in each cohort accounted for by private sector employment in 1994–96 (average) is as follows: 1926–1935: 0.21; 1936–1945: 0.24; 1946–1955: 0.30; 1956–1965: 0.34; 1966–1975: 0.48; 1976–1985: 0.76. We do not report results in the table for cells with fewer than 100 observations.

²³ As Juhn, Murphy and Pierce also note (p. 425), while the average (across cohorts) change in inequality combines age and time effects, taking the *change* in this measure over time causes age effects to drop out. Thus, the increase in the growth rate of inequality we observe between the 1985–1988 period and the 1988–1992 period must be due to time effects. The same is true for the decrease in the growth rate of inequality we observe between 1988–1992 and 1992–1996.

²⁴ The residuals are from annual OLS regressions of log earnings on a constant, four education dummies, experience and its square, and dummies for gender, urban residence, and employment in the private sector. These are identical to the specifications examined in greater detail in the regression analysis of Section 5.

In order to compare trends in residual earnings inequality across groups, we regressed the squared earnings residuals on the same covariates used in the earnings regressions, plus a time trend and an interaction of this trend with worker characteristics. These interaction terms capture the change over time in within-group inequality controlling for all other worker attributes. The only significantly positive trend interaction with the education dummies was that for college-educated workers, again confirming the relatively higher increase in inequality within this group, even after controlling for other characteristics.²⁵ Consistent with the earlier results, we also found significant positive trend interactions with the male, private sector and urban residence dummies.

To summarize, within-group inequality rose significantly among college-educated workers, males, private sector workers and urban workers. Increases in inequality seem to have been quite similar within broadly defined experience groups.

4.6. Earnings inequality among men and women

Next, we examine the evolution of earnings inequality among men and women. Most of the literature on earnings inequality in industrialized countries has focused on male wage inequality. This omission is largely driven by the fact that, in such countries, women typically have much lower employment rates than men, and are more likely to enter and exit from employment over time, introducing selection problems. But in Poland, women have traditionally had employment rates nearly as high as those for men, and in our sample they make up 45% to 48% of employed workers. Thus, it is of considerable interest to analyze these two groups separately, and also to examine their respective impacts on the overall earnings distribution.

The bottom right panels of [Table 2](#) show percentile differentials separately for men and women. The increase in earnings inequality was much greater for men than women. For instance, from 1988 to 1996, the increase in the 90–10 differential is 0.21 for men and 0.15 for women. Similarly, the increase in the 75–25 differential is 0.12 for men and 0.07 for women.

It is worth noting that the increase in male earnings inequality is greater than the increase in overall inequality. As [Fortin and Lemieux \(2000\)](#) discuss in the U.S. context, changes in overall earnings inequality depend both on how inequality changes for males and females separately, as well as on the distance between the male and female earnings distributions. To examine how the gap between the distributions contributed to changes in overall inequality, [Fig. 3](#) presents employment-share weighted kernel density estimates for earnings of men and women for the years 1988, 1992, 1994 and 1996. Note that, throughout the period 1988–1996, the female earnings density is clearly to the left of the male density. It is also hard to discern any tendency for the two densities to converge over time.

²⁵ The time trend is 0.063 for the college-educated group (standard error=0.004). It is 0.009 for the some-college group (not significant), 0.032 for the high school group, 0.008 for the vocational training group, and 0.015 for the primary school group. We also ran these types of regressions separately for public and private sector workers, and, in each sector, the time trend was much larger for the college educated group. The estimated coefficient on the interaction of time with the college degree dummy was 0.056 in the public sector and 0.101 in the private sector.

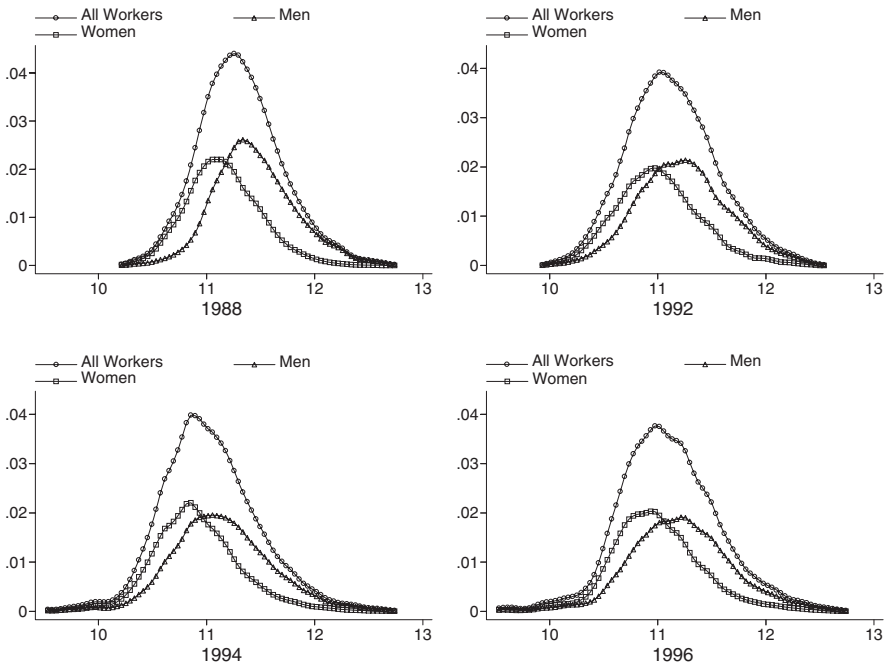


Fig. 3. Weighted kernel density estimates: by gender.

Fig. 1 (right panel) shows that the median wage differential between men and women fell from about 0.3 to about 0.2 during the early stages of transition, but then rebounded and, after 1992, has stabilized at around 25%. Although the wage premium for men is higher in the private sector in the early years of transition, it, too, levels off at about 25% after 1992.

In contrast, Hunt (2002) reports that German unification resulted in a 10 percentage point decline in the gender wage gap in the former East Germany but notes that this was accompanied by a much greater decline in female employment than in male employment. She also finds that low earners were more likely to leave employment after unification and that such persons were disproportionately female. In Poland, by contrast, the share of women in total employment actually increased after the transition (see Table 1). Thus, there appear to be very different forces at work in Poland in terms of influences on the gender wage gap. We leave a more detailed analysis of this issue for future research.

5. Regression analysis of the structure of earnings

5.1. Human capital earnings functions

In this section, we conduct a regression analysis of the structure of earnings in order to gain more insight into the sources of changing inequality among workers during the transition. In Table 6, we present estimates of standard human capital earnings functions

Table 6
OLS earnings regressions

	1986	1988	1989	1990	1991	1992	1994	1995	1996
College degree	0.368* (0.009)	0.383* (0.009)	0.430* (0.010)	0.497* (0.011)	0.550* (0.010)	0.527* (0.018)	0.636* (0.010)	0.632* (0.010)	0.683* (0.010)
Some college	0.225* (0.027)	0.229* (0.025)	0.270* (0.032)	0.394* (0.031)	0.408* (0.034)	0.334* (0.058)	0.349* (0.014)	0.348* (0.015)	0.386* (0.014)
High school	0.180* (0.007)	0.205* (0.007)	0.209* (0.008)	0.238* (0.008)	0.276* (0.008)	0.264* (0.014)	0.306* (0.008)	0.308* (0.008)	0.343* (0.008)
Basic vocational training	0.132* (0.007)	0.135* (0.007)	0.115* (0.007)	0.106* (0.008)	0.112* (0.008)	0.120* (0.014)	0.149* (0.008)	0.151* (0.008)	0.165* (0.008)
Experience	0.035* (0.001)	0.032* (0.001)	0.036* (0.001)	0.030* (0.001)	0.030* (0.001)	0.029* (0.002)	0.030* (0.001)	0.030* (0.001)	0.030* (0.001)
Experience sqrd./100	-0.062* (0.002)	-0.056* (0.002)	-0.065* (0.002)	-0.052* (0.002)	-0.050* (0.002)	-0.048* (0.004)	-0.053* (0.002)	-0.052* (0.002)	-0.051* (0.002)
Male	0.326* (0.005)	0.301* (0.005)	0.280* (0.005)	0.247* (0.005)	0.243* (0.005)	0.260* (0.009)	0.293* (0.005)	0.302* (0.005)	0.299* (0.005)
Urban	0.052* (0.005)	0.071* (0.005)	0.042* (0.005)	0.046* (0.005)	0.058* (0.005)	0.095* (0.011)	0.126* (0.005)	0.115* (0.005)	0.115* (0.005)
Private enterprise	0.075* (0.014)	0.059* (0.012)	0.067* (0.111)	0.082* (0.010)	0.130* (0.008)	0.090* (0.012)
Small private firms	-0.087* (0.006)	-0.110* (0.005)	-0.102* (0.005)
Large private firms	0.180* (0.021)	0.189* (0.021)	0.183* (0.018)
Large mixed-ownership firms	0.075* (0.022)	0.087* (0.023)	0.106* (0.025)
Adjusted R ²	0.301	0.272	0.235	0.220	0.264	0.242	0.297	0.302	0.324
Number of observations	19,666	21,114	20,787	20,685	19,025	7,259	26,095	26,316	26,165

The sample includes employed workers between the ages of 18 and 60, who report labor income as their primary source of income. The dependent variable is log quarterly real labor income for 1985–1992 and log monthly real income for 1994–1996. Robust standard errors are reported in parentheses. An asterisk indicates statistical significance at the 5 percent level.

(see Willis, 1986) for a selected set of years during the 1985–1996 period.²⁶ The coefficients on the education category dummies show that education premiums have increased substantially during the transition. For instance, the college degree coefficient increases from 0.383 in 1988 to 0.527 in 1992 and further to 0.683 in 1996. This implies that the earnings premium for a college degree relative to a primary school education (the omitted group) was approximately 47% in 1987, 69% in 1992, and 98% in 1996. The high school premiums relative to primary school for the same three years are 23, 30 and 41 percent, respectively. These results imply a widening of the college–high school premium as well. By 1996, the high school and college premiums in Poland were toward the high end of estimates typically obtained using data from Western countries.

Our findings of a sharp increase in education premiums after the transition are consistent with those of Gorecki (1994), based on his examination of aggregate Polish wage data, and of authors who have examined the wage structure in other transition economies. For instance, Ham et al. (1995) examine surveys conducted by the Federal Ministry of Labor in Czechoslovakia in 1988 and 1991. They find that the wage gap between university and elementary school graduates increased from 58% in 1988 to 63% in 1991. Brainerd (1998) examines survey data for Russia and reports that, from 1991 to 1994, the marginal return to a year of education rose from 3.1% to 6.7% for men and from 5.4% to 9.6% for women.

One interesting question is whether academic qualifications acquired in the communist era have the same value in the labor market as more recently acquired qualifications. Unfortunately, it is difficult to answer this question decisively with our sample, since it ends in 1996, seven years after the big bang. As a first pass, we constructed a dummy variable for individuals who, based on their age and imputed years of education, attained their highest degree on or after 1992. This dummy, and its interactions with the education dummies, did not show any strong evidence that recent degrees command a premium relative to older degrees.²⁷

Since (potential) labor market experience (age-years of education-6) enters the regressions as a quadratic, the returns to experience need to be evaluated at specific levels of experience. The OLS estimates imply a decline in the returns to experience in the early years of the transition, with a subsequent rebound.²⁸ For instance, for workers with 25 years of potential labor market experience, the return to an additional year of experience is 0.41% in 1988, drops to 0.36% in 1990 and then rises to 0.48% by 1996. For younger workers with 5 years of potential experience, this return drops from 2.64% in 1988 to 2.46% in 1990 and then partially recovers to 2.52% by 1996. Note that, in any case, these returns to experience are much smaller than those typically found in Western data sets.

Consistent with the evidence from the previous section, the earnings premium for men relative to women drops by about 5% points in the early years of transition, but then rises

²⁶ We continue to use the adjusted data for 1994–1996. However, whether we use adjusted or unadjusted data makes no difference to the point estimates of the regression coefficients since the OLS regressions are essentially identical to those used in the first stage regressions for the adjustment procedure.

²⁷ Certain Polish observers have told us that major curriculum changes in the post-Communist era have taken place mainly at the university level. There also appears to be some debate about whether the level of rigor in school and university programs has improved in the post-communist era, with some actually arguing the opposite.

²⁸ Lehmann and Wadsworth (2000) report similar results for Poland as well as Russia.

again and stabilizes around 30% after 1992. The earnings premium for workers in urban areas relative to those in rural areas, on the other hand, almost doubles by the mid-1990s relative to pre-transition levels.

Until 1992, the HBS asked if the respondent worked for a privately owned firm. In the 1994–1996 surveys, this question was refined, and we can determine if the respondent works for one of three types of privately owned firms: small firms (including the self-employed); large privately held firms; and large mixed-ownership firms with majority ownership by the private sector. So, in the later years, we include separate dummies for these three categories.

In the years up until 1992, the private sector dummy coefficient is always significantly positive, and generally around 0.08. Recall from Fig. 1 that the unconditional private sector earnings premium was around 10% from 1985 to 1992, but that this premium turned sharply negative in 1994–1996. Our results for 1994–1996 indicate that earnings of workers in small private firms (including the self employed) were about 10% below those of similar workers (controlling for observed education, experience, etc.) in the public sector. But workers in large private firms (including foreign-owned firms) earn approximately 20% more than observationally similar workers in the public sector, while people who work for large firms with mixed but majority private ownership earn about 11% more. A dummy for employment in mixed-ownership firms with majority state ownership was small and not significant.²⁹

We also ran regressions separately for workers in the public and private sectors. Some of the main results, and a comparison with the results for the full sample, are summarized in Fig. 4. Two aspects of the results are worth noting. First, education premiums rose much more sharply in the private sector during the transition. In both sectors, the college and high school premiums were about 40% and 20%, respectively, before the transition. By 1996, these premiums had risen to 60% and 30% in the public sector, but in the private sector they had risen to about 90% and 40%, respectively.

Second, we note that, for older workers (25 years of potential labor market experience) in the private sector, the return to an additional year of experience was close to zero before the transition. After the transition, this return begins to rise and, by 1996, is close to 0.6%, even higher than in the public sector (0.4%). We conjecture that this reflects selection effects and the fact that experienced workers who were able to obtain private sector employment after the transition have strong unobserved attributes. For younger workers (5 years of potential labor market experience), the returns to an additional year of experience are in the range of 2.5–3.0 percent in both sectors before the transition. By 1996, this return declines to 2.3% in the private sector and rises marginally to 2.7% in the public sector. A possible explanation for this pattern is that the relative supply of younger workers looking for private sector jobs rose sharply during the transition, thereby driving down the returns to experience (at low levels of experience) in the private sector. This is consistent with our earlier

²⁹ We also ran the regressions for 1994–96 with just an overall private sector dummy. The estimated coefficient was in the range of -0.08 to -0.10 in 1994–1996. This is similar to, although slightly smaller than, the sharp reversal of the (unconditional) private sector earnings premiums after 1992 shown in Fig. 2.

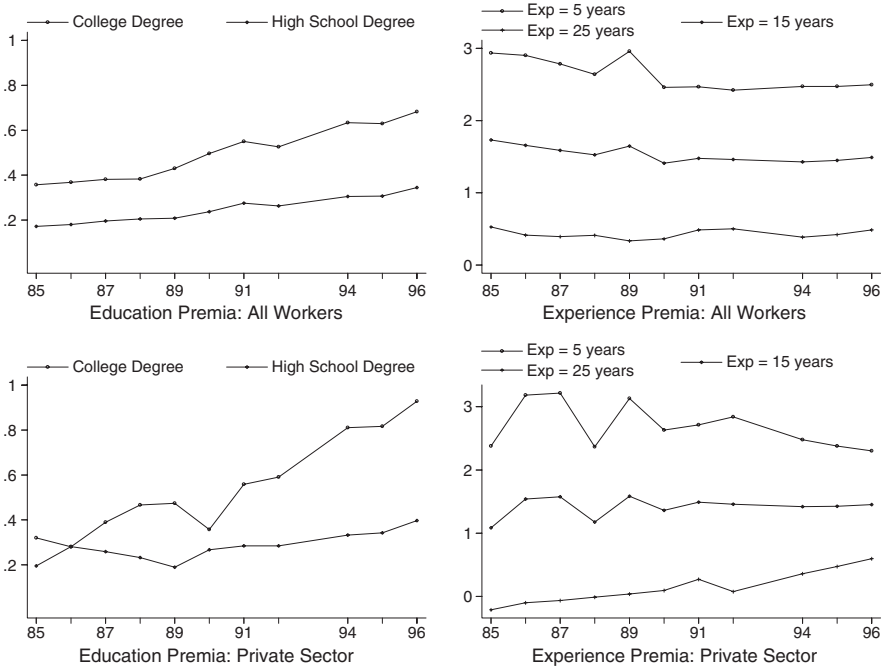


Fig. 4. Education and experience premiums.

finding that the most recent birth cohort has a very high level of private sector employment.

We also computed the wage regressions separately for men and women. To conserve space, we do not report those results in detail. One intriguing finding was that, before the transition, education premiums were substantially higher for women than for men. For instance, in 1988, the college and high school premiums (relative to a primary school degree) for women were 58% and 30%, respectively, compared to 38% and 16%, respectively, for men. By 1996, these premiums for both men and women had converged to about 95% and 40%. Thus, although both men and women experienced substantial increases in education premiums during the transition, these increases were much greater for men.

A possible explanation for this pattern, given the greater increases in education premiums in the private sector noted earlier, is that men have disproportionately benefited from employment growth in the private sector. But a cursory examination of the data strongly refutes this hypothesis. Although the share of women and men in public sector employment has remained roughly equal over our entire sample, the share of private sector employment accounted for by women in fact rose from about 20% before the transition to about 40% by 1996.

5.2. Quantile regressions

In this section, we use quantile regressions to obtain a more complete picture of changing returns to education and experience, and to characterize in a parsimonious way

the changes in the entire conditional distribution of income (see Buchinsky, 1994). We ran quantile regressions of log real quarterly (monthly for 1994–1996) labor income on the same set of regressors as in the OLS regressions. Regressions were run for the 0.10, 0.25, 0.50, 0.75 and 0.90 quantiles.

Table 7 reports the college and high school premiums (relative to primary school) from the quantile regressions. Note that the college premium jumps sharply from 1988 to 1996 at all quantile points. For example, median earnings were approximately 48% ($\exp(0.39)$) higher for college graduates than primary school graduates in 1988, and this premium increases to 95% ($\exp(0.67)$) by 1996. However, the increases in college premiums are greater at higher quantiles. Thus, by 1996, the college premium is considerably greater at both the 0.75 and 0.90 quantiles than it is at the median. At the 0.90 quantile, it is as high as 120%. The high school premium also rises substantially from 1988 to 1996 at all quantile points. For example, median income was approximately 20% higher for high school graduates compared to primary school graduates in 1988, and this premium increases to 43% in 1996. The levels of the high school premium, and its behavior during the transition years, are quite similar across all quantiles.

When we ran the quantile regressions limiting the sample to private sector workers, we found greater increases in the levels of college premiums by 1996 at all quantile points. For instance, in 1996 the college premium was about 150% at the median of the distribution and over 170% at the 0.90 quantile in the private sector. The high school premium is also higher in the private sector, although only marginally. By 1996, the pattern of distribution across quantiles of college premiums (higher at the upper quantiles) and high school premiums (similar across quantiles) in the private sector is quite similar to that in the full economy.

Overall, the quantile regressions reveal a rather interesting pattern in terms of how education premiums changed during the transition. While the changes in the high school premium were similar at different points of the distribution, the growth of the college premium was markedly greater at higher quantiles.

Table 8 shows returns to experience, defined as the derivative of log real labor income with respect to labor market experience. We report the derivatives evaluated at 5, 15 and 25 years of experience. The numbers in Table 8 are in percentage terms. So, for example, in 1986, for workers with 5 years of experience, an additional year of experience raises earnings by about 2.9% at the (conditional) median of the wage distribution. For those with 15 years of experience, the return to experience at the median in 1986 is smaller (1.67%) and, for those with 25 years of experience, it is even smaller (0.42%). Results are similar at other quantile points. These experience effects are much smaller than those found in Western data sets.³⁰

In the immediate aftermath of the big bang (i.e., 1990) returns to experience dropped sharply relative to their pre-transition levels. This is true at almost all quantile points for workers with 5 or 15 years of experience. In the 1992–1996 period, there is some recovery in experience returns, especially at higher experience levels. In 1996, the returns to experience at the 25-year experience level range from 0.59% at the 0.10

³⁰ For instance, Buchinsky (1994) finds returns to experience for U.S. workers with 15 years of experience to be about 2.9 to 3.0 percent at all quantile points in 1985, nearly double the figure we obtain for Poland in 1985.

Table 7
Education premiums: quantile regressions

	Quantile=0.10		Quantile=0.25		Quantile=0.50		Quantile=0.75		Quantile=0.90	
	COL	HS	COL	HS	COL	HS	COL	HS	COL	HS
1986	0.41 (0.01)	0.23 (0.01)	0.39 (0.01)	0.21 (0.01)	0.38 (0.01)	0.18 (0.01)	0.38 (0.01)	0.17 (0.01)	0.35 (0.02)	0.16 (0.01)
1988	0.39 (0.01)	0.24 (0.01)	0.38 (0.01)	0.22 (0.01)	0.39 (0.01)	0.20 (0.01)	0.39 (0.01)	0.20 (0.01)	0.39 (0.02)	0.20 (0.01)
1989	0.42 (0.02)	0.22 (0.01)	0.43 (0.01)	0.21 (0.01)	0.45 (0.01)	0.21 (0.01)	0.45 (0.02)	0.20 (0.01)	0.43 (0.02)	0.22 (0.02)
1990	0.49 (0.02)	0.26 (0.01)	0.50 (0.01)	0.25 (0.01)	0.51 (0.02)	0.23 (0.01)	0.53 (0.02)	0.23 (0.01)	0.53 (0.02)	0.24 (0.01)
1991	0.53 (0.02)	0.26 (0.01)	0.53 (0.01)	0.26 (0.01)	0.56 (0.02)	0.28 (0.01)	0.59 (0.02)	0.29 (0.01)	0.62 (0.02)	0.33 (0.01)
1992	0.50 (0.03)	0.24 (0.03)	0.51 (0.02)	0.25 (0.01)	0.52 (0.03)	0.26 (0.02)	0.59 (0.03)	0.30 (0.02)	0.58 (0.03)	0.31 (0.03)
1994	0.61 (0.01)	0.33 (0.01)	0.60 (0.01)	0.32 (0.01)	0.63 (0.01)	0.32 (0.01)	0.69 (0.02)	0.33 (0.01)	0.75 (0.02)	0.36 (0.01)
1995	0.62 (0.02)	0.34 (0.02)	0.61 (0.01)	0.33 (0.01)	0.65 (0.01)	0.33 (0.01)	0.69 (0.01)	0.33 (0.01)	0.73 (0.02)	0.34 (0.01)
1996	0.69 (0.01)	0.39 (0.01)	0.65 (0.01)	0.36 (0.01)	0.67 (0.01)	0.36 (0.01)	0.75 (0.02)	0.37 (0.01)	0.79 (0.02)	0.37 (0.02)

COL—college degree; HS—high school degree. The full set of regressors included four education dummies, experience and its square, and dummies for gender, urban residence and employment in the private sector. The dependent variable is log real quarterly (monthly for 1994–1996) labor income. The excluded education dummy is for primary school degree. Hence, the coefficients reported above are interpretable as the income premiums, in percent (multiply by 100), for workers with a college or high school degree, respectively, relative to workers who have only a primary school degree. Bootstrapped robust standard errors are reported in parentheses below the coefficient estimates.

Table 8
Returns to experience: quantile regressions

Experience (in years):	5 years					15 years					25 years				
	0.10	0.25	0.50	0.75	0.90	0.10	0.25	0.50	0.75	0.90	0.10	0.25	0.50	0.75	0.90
Quantile point:															
1986	2.79 (0.10)	2.86 (0.05)	2.92 (0.09)	3.03 (0.11)	3.01 (0.15)	1.67 (0.04)	1.68 (0.03)	1.67 (0.04)	1.71 (0.06)	1.70 (0.08)	0.55 (0.05)	0.50 (0.02)	0.42 (0.02)	0.39 (0.04)	0.39 (0.07)
1988	2.40 (0.11)	2.45 (0.09)	2.58 (0.08)	2.76 (0.07)	2.77 (0.20)	1.45 (0.06)	1.45 (0.05)	1.52 (0.04)	1.59 (0.04)	1.56 (0.11)	0.49 (0.03)	0.46 (0.03)	0.46 (0.03)	0.41 (0.03)	0.34 (0.04)
1989	2.93 (0.12)	2.83 (0.06)	2.91 (0.07)	3.02 (0.13)	3.07 (0.14)	1.66 (0.07)	1.60 (0.03)	1.64 (0.04)	1.68 (0.07)	1.68 (0.07)	0.39 (0.05)	0.37 (0.05)	0.37 (0.03)	0.35 (0.05)	0.29 (0.07)
1990	2.55 (0.15)	2.49 (0.11)	2.36 (0.08)	2.51 (0.12)	2.37 (0.15)	1.47 (0.08)	1.45 (0.06)	1.38 (0.05)	1.42 (0.06)	1.35 (0.08)	0.40 (0.05)	0.40 (0.04)	0.39 (0.03)	0.37 (0.04)	0.33 (0.06)
1991	2.35 (0.13)	2.37 (0.11)	2.37 (0.13)	2.50 (0.11)	2.72 (0.18)	1.41 (0.07)	1.43 (0.06)	1.45 (0.06)	1.52 (0.06)	1.61 (0.09)	0.48 (0.04)	0.49 (0.04)	0.54 (0.05)	0.53 (0.04)	0.49 (0.05)
1992	2.19 (0.27)	2.23 (0.11)	2.47 (0.14)	2.42 (0.23)	2.62 (0.31)	1.40 (0.13)	1.39 (0.06)	1.50 (0.08)	1.50 (0.12)	1.54 (0.17)	0.62 (0.10)	0.54 (0.08)	0.52 (0.08)	0.58 (0.08)	0.47 (0.08)
1994	2.31 (0.09)	2.26 (0.09)	2.44 (0.11)	2.56 (0.10)	2.64 (0.17)	1.41 (0.04)	1.37 (0.04)	1.44 (0.05)	1.45 (0.05)	1.46 (0.09)	0.51 (0.05)	0.47 (0.05)	0.43 (0.03)	0.34 (0.04)	0.28 (0.10)
1995	2.06 (0.08)	2.17 (0.10)	2.47 (0.08)	2.84 (0.13)	3.17 (0.12)	1.31 (0.04)	1.34 (0.05)	1.48 (0.04)	1.64 (0.06)	1.73 (0.06)	0.55 (0.05)	0.50 (0.03)	0.49 (0.03)	0.44 (0.05)	0.29 (0.06)
1996	2.18 (0.10)	2.25 (0.06)	2.49 (0.09)	2.73 (0.11)	2.96 (0.16)	1.39 (0.04)	1.39 (0.03)	1.50 (0.05)	1.58 (0.05)	1.69 (0.07)	0.59 (0.05)	0.54 (0.02)	0.51 (0.04)	0.43 (0.04)	0.41 (0.05)

The numbers reported above are derivatives of log real quarterly (monthly for 1994–1996) labor income with respect to the experience level of the worker, evaluated at the indicated experience levels. The results are expressed in percentage terms. Standard errors for these derivatives, based on bootstrapped robust standard errors for the regression coefficients, are reported in parentheses. The full set of regressors included four education dummies, experience and its square, and dummies for gender, urban residence and employment in the private sector.

quantile down to 0.41% at the 0.90 quantile. These figures are much higher than those for 1990, especially at the lower quantiles. After 1991, experience premiums are relatively stable for workers with 15 years of experience but continue to decline for younger workers with only 5 years of experience. Thus, somewhat surprisingly, recent labor market entrants are the only group for whom experience returns are systematically lower in 1996 than in 1987.

In earlier work (Keane and Prasad, 2002), we noted the marked increase in the relative generosity of pensions in 1991–1992 and the consequent surge in the pension rolls in those years. Indeed, in our sample, the share of total employment accounted for by workers in the 51–60 age group fell from about 13% in 1989–1990 to 11% in 1991–1992 and, further, to 9% by 1996. In other words, it is possible that self-selection of workers with weaker unobserved attributes into retirement could account for the recovery in experience premiums observed in the mid-1990s at higher experience levels.

In summary, a key finding from this section is the very different behavior of education premiums and experience premiums during the transition. Starting from relatively low levels in the pre-transition period, education premiums had, by 1996, risen to a level in line with, if not somewhat greater than, those typically found in Western industrial economies. In contrast, experience premiums, which were small to begin with, generally became even smaller. Thus, the divergence in returns to experience compared to countries like the U.S. became even larger than before the transition.

5.3. *The industry structure of earnings and employment*

In this section, we analyze changes in the industry structure of wages, and the allocation of workers of different skill types across industries. As the starting point for this analysis, we re-ran the log earnings regressions from Table 6, but this time including dummies for 13 industry classifications. The estimates of the industry dummy coefficients for selected years are reported in Table 9. The table also reports the share of workers employed in each industry.

We find that earnings differentials across industries are very stable during 1985–1988. Every industry has a negative differential (usually large) relative to manufacturing. This means that workers who were equivalent in terms of education, experience, gender, urban/rural residence and private/state sector employment earned more in manufacturing industries than in other industries.

However, in the aftermath of the big bang, industry differentials changed very quickly. By 1991, earnings differentials between almost all industries and Manufacturing narrowed considerably, and even turned positive in Finance and Insurance, Public Administration, and Real Estate. Clearly, even conditional on education (controlled for in the regression), relative wages rose in those industries that one would typically classify as white collar and/or requiring more cognitive skills.

Between 1991 and 1996, earnings differentials again changed substantially. Those between the relatively white collar industries and Manufacturing were uniformly much smaller than in 1988, and the differential in Science and Technology also turned

Table 9
Industry wage differentials and employment shares

Industry:	1985		1988		1991		1996	
	Earnings	Employment	Earnings	Employment	Earnings	Employment	Earnings	Employment
Manufacturing	–	39.2	–	37.2	–	35.8	–	30.6
Construction	–0.059	8.6	–0.064	8.7	–0.020	8.1	–0.065	8.0
Agriculture, Forestry, Fishing	–0.038	9.0	–0.070	11.7	–0.062	9.8	–0.203	4.7
Transport and Communication	–0.098	8.7	–0.122	8.5	–0.028	8.6	–0.031	8.5
Trade	–0.140	8.9	–0.143	9.6	–0.058	9.1	–0.051	11.5
Real Estate	–0.132	4.2	–0.149	3.7	0.004	3.9	0.001	3.0
Science and Technology	–0.138	1.2	–0.139	1	–0.031	0.9	0.052	0.7
Education and Culture	–0.200	8.6	–0.263	8.6	–0.058	9.7	–0.154	9.7
Health and Hygiene	–0.193	6.0	–0.199	5.8	–0.037	6.9	–0.101	8.5
Recreation and Tourism	–0.160	0.6	–0.158	0.7	–0.075	0.6	–0.069	1.2
Other Services	–0.155	0.9	–0.135	1.2	–0.018	1.3	–0.126	3.8
Public Administration	–0.103	3.0	–0.120	2.5	0.099	3.9	–0.013	7.3
Finance and Insurance	–0.031	1.0	–0.096	0.9	0.213	1.4	0.086	2.5

Bold figures for 1991 indicate industries whose share increased from 1989–1996, and where earnings in 1991 substantially overshot their ultimate 1996 levels. Bold figures for 1996 indicate industries where wages rose most substantially relative to manufacturing during the transition. All industry earnings differentials are from log earnings equations that control for the same covariates as in Table 6. Manufacturing is the omitted category.

positive. The set of industry dummy coefficients is quite stable between 1995 and 1996, suggesting that the industry wage structure has settled down to a new equilibrium by the end of our sample period. The industries where earnings rose most substantially relative to manufacturing during the transition are highlighted in bold in 1996. These are Finance and Insurance, Science and Technology, Real Estate and Public Administration.

In 1990–1991 we see a clear phenomenon whereby industry relative earnings overshot their ultimate 1996 levels in 5 of the 7 industries that expanded during the transition (i.e., Finance and Insurance, Public Administration, Other Services, Health and Hygiene, Education and Culture). The earnings differentials for these 5 industries are highlighted in bold for 1991. The only exceptions to this pattern are Trade and Recreation and Tourism, which expanded but where earnings differentials were similar in 1991 and 1996. The 7 expanding industries are all ones that would typically be classified as relatively white collar.

Next, in Table 10, we report (for selected years) the education coefficients from log earnings equations with and without industry dummies. The coefficients for the specifications without industry dummies are simply reproduced from Table 6. The

Table 10
Education coefficients with and without industry dummies

	1988		1991		1996	
College	.383	.475	.550	.553	.683	.705
Some college	.229	.306	.408	.417	.386	.402
High school	.205	.224	.276	.264	.343	.334
Vocational training	.135	.128	.112	.107	.165	.153
Industry dummies		YES		YES		YES
Adj. R^2	.272	.311	.264	.275	.324	.341

The education coefficients are taken from the log earnings regressions reported in Tables 6 and 10. Primary education is the omitted category.

following interesting pattern emerges: in the pre-transition period, the College and High School earnings premiums (relative to Primary school) are substantially increased (by 8 to 9 points) when industry is controlled for. This is consistent with a pattern whereby better educated workers are sorted into “low wage” industries (i.e., those with negative earnings differentials in Table 9). But, in 1990–1991, this pattern completely vanishes. Education premiums are almost unchanged when industry is controlled for. This is consistent with the substantial increase in relative wages in several white collar industries that tend to employ more educated workers (which we saw in Table 9). Note also that the contribution to R^2 from including industry dummies is much greater in the pre-transition period.

Next, we examine the sorting of workers (by skill level) across industries. To do this, we grouped industries into three categories. Group 1 includes Manufacturing, Construction and Agriculture. Group 2 includes Transport and Communication, Trade, Real Estate, Recreation and Tourism, and Other Services. Group 3 includes Science and Technology, Education and Culture, Health, Public Administration, and Finance and Insurance. In Table 11, panel A, we report the allocation of workers, by education level, across these three industry groups, in 1988 and 1996.

Notice that, in 1988, 69% of workers with a vocational training, and 65% of workers with primary school training, were employed in Group 1. In contrast, 56% of college educated workers were employed in Group 3. These figures are consistent with a characterization of industries in Group 1 as being typically blue collar/low skilled, and of the Group 3 industries as being typically white collar or high skilled. The Group 2 industries are in the middle, so we refer to them as medium skilled.

One striking pattern evident in Table 11 panel A is the shift of college educated workers out of Group 1 industries. The percentage of college educated workers employed in industry Group 1 drops from 34% in 1988 to only 19% in 1996. Notably, workers of all types exit Group 1, and the fraction of all workers employed in group 1 falls from 58% to 43%.

In panel B of Table 11, we describe the composition of industries by education level. The most notable pattern here is skill upgrading: In all three industry groups, the percentage of workers with only a primary school education falls. In the employed workforce as a whole, the percentage of workers with only a primary education fell from 21.2% in 1988 to 11.9% in 1996 (a 44% decline). This pattern arises for two reasons. First, primarily due to cohort effects, the average education level of the 18–60 year old population increased from 1988 to 1996. The fraction with only primary education fell

Table 11
The changing allocation of workers across industries by education level

A. Allocation of education groups across industries								
	Group 1		Group 2		Group 3		1988	1996
	1988	1996	1988	1996	1988	1996		
College	33.7	19.1	10.0	16.5	56.3	64.4		
High school	46.0	34.8	27.2	29.8	26.8	35.4		
Vocational	68.8	55.0	24.5	30.6	6.8	14.4		
Primary	65.3	56.9	22.8	25.7	11.9	17.4		
All	57.6	43.3	23.6	28.0	18.8	28.7		

B. Employment distribution across industries by education level								
	Group 1		Group 2		Group 3		All industries	
	1988	1996	1988	1996	1988	1996	1988	1996
College	5.4	5.3	3.9	7.1	27.7	27.1	9.3	12.1
High school	25.2	30.2	36.3	40.0	45.1	46.4	31.6	37.6
Vocational	45.3	48.8	39.3	42.0	13.7	19.3	37.9	38.5
Primary	24.1	15.7	20.5	11.0	13.5	7.3	21.2	11.9

Industry groups are:

- 1—Manufacturing, construction, agriculture, forestry, fishing.
- 2—Transport and communications, trade, real estate, recreation and tourism, other services.
- 3—Science and technology, education and culture, health, public administration, finance and insurance.

from 36.0% to 25.8% (a 28% decline). Accentuating this, generous pension benefits led to substantial early retirement among less-educated workers in the 51–60 age group in 1991–1992. Consequently, among people with just a primary education, the fraction of those who were employed declined by 39% from 1988 to 1996.

To summarize, the early phase of the transition was characterized by dramatic shifts of relative wages, favoring higher skilled industries at the expense of blue collar industries like manufacturing. This induced a substantial reallocation of workers out of manufacturing and into higher skilled industries. The realignment of industry relative wages during the transition greatly reduces the explanatory power of industry dummies in earnings functions.

6. Discussion

As we noted in the introduction, we cannot discern the fundamental causes of the changing structure of earnings during the Polish transition based on the descriptive statistics or “stylized facts” we have developed here. However, we can ask whether various models of transition and/or the evolution of earnings inequality that have been developed in the literature appear consistent with the stylized facts, and, if not, what sorts of amendments to these models might help them match the data.

In fact, existing models of transition do not capture key aspects of the changing structure of earnings we see in Poland. For instance, [Aghion and Commander \(1999\)](#) and

Commander and Tolstopiatenko (1998) present models of transition with a state sector and a private sector. In the state sector, firms operate under a zero-profit constraint with no capital accumulation (insiders extract all surplus) and wages equal average product. Private sector firms behave competitively. The state sector is less efficient (captured by a lower constant in the Cobb–Douglas production function). State firms face a probability of closure, and can choose to restructure, in which case they shed excess labor and become like private firms. There is unemployment if hiring by private firms falls short of exits from state firms. Initial inequality within each sector is set exogenously in the model, and the authors assume a higher level in the private sector.

When this model is simulated, higher inequality arises during the transition for two main reasons: 1) workers shift from the relatively low inequality state sector to the higher inequality private sector, and 2) mean wages are higher in the private sector (due to greater productivity). The latter feature leads to a Kuznets curve (i.e., an inverse U-shape for inequality) because, with a mean wage differential between the two sectors, the contribution of between sector inequality to total inequality peaks when the allocation of workers between the two sectors is equal.

As our analysis reveals, this type of model misses several key features of the data. First, mean wage differences between sectors are not an important factor driving changes in inequality (and, furthermore, mean wages are actually lower in the private sector than in the public sector). Second, inequality has grown within both sectors over time, and this is the main factor driving changes in inequality. Third, inequality has not followed a Kuznets curve but has trended upward until reaching a plateau. Fourth, the model does not account for the greater increase of inequality among the more educated, which also happens to be the group that has had the *least* reallocation into the private sector.

This type of model could be modified to explain rising inequality of earnings in the state sector if it accounted for restructuring of SOEs in the absence of privatization—the phenomenon that Pinto et al. (1993), Commander and Dhar (1998) and others suggest is important. Clearly, as the fraction of firms engaged in competitive wage setting grows (both through increases in the size of the private sector and restructuring of SOEs), the relative demand for skilled labor will increase in the economy as a whole. Through this mechanism, inequality will also grow in the private sector during the transition.³¹

Assuming the state sector has moved closer to competitive wage setting, another source of increasing inequality within that sector may have been price liberalization and opening to trade. For example, energy prices were held artificially low prior to 1990, and oil could be imported cheaply from the U.S.S.R. As Keane and Prasad (1996) note, there is a large body of empirical work suggesting that energy is complementary with unskilled labor and substitutable with both capital and skilled labor in manufacturing production. Thus, an increase in the price of energy would increase the demand for skilled labor relative to the unskilled labor. This is the same mechanism whereby a falling price of capital increases skill premiums in Krusell, Ohanian, Rios-Rull and Violante (2000).

³¹ Of course, another mechanism driving increased inequality within the state sector may simply be the need to compete for labor with the growing private sector. Increased labor mobility across sectors, industries and firms would naturally be expected to raise within-group inequality as search gives rise to an equilibrium wage distribution.

In addition, the transition can be interpreted as a period of technical change, not in the sense that the technical frontier has shifted out, but rather in the sense that SOEs have reorganized to achieve greater technical efficiency, and modified production processes to reflect market prices of inputs (in particular, by eliminating labor hoarding). Much of the existing capital was presumably rendered obsolete. There are a number of models where more educated workers are better able to adapt to rapid changes in production processes (see, e.g., [Acemoglu, 1998](#); [Caselli, 1999](#)), and this mechanism may help explain rising returns to skill.

Another notable aspect of the transition is that returns to general human capital, as measured by education premiums, rose markedly, while the returns to experience remained small. Low returns to experience appear consistent with the notion of rapid obsolescence of firm- or industry-specific skills during a period of rapid technological change and industrial restructuring (see [Svejnar, 1996](#)).

A rapid pace of change may also help explain rising within-group inequality. For instance, [Violante \(2002\)](#) proposes a model in which technology is embodied in capital of different vintages. Given frictions in the matching of workers with new technology, acceleration of technical progress (i.e., updating of capital) increases wage dispersion among identical workers, thus increasing within-group inequality. Similarly, [Aghion and Commander \(1999\)](#) argue that transition can be viewed as generating a technology gap between the old state sector and the new entrepreneurial sectors. Growth in within-group inequality might then be explained by [Violante's \(2002\)](#) matching framework.

Clearly, however, this framework needs to be modified to account for the fact that increases in within-group inequality in Poland occurred primarily for highly educated workers. One possible approach is the model of [Galor and Moav \(2000\)](#), where rising returns to education increase within-group inequality among the highly educated through a compositional effect. They model human capital as being a function of innate ability as well as formal education. In their model, it is costlier for low-ability individuals to acquire education. Technological progress that increases the returns to education without affecting the return to innate ability leads some lower-ability individuals to invest in education, thereby widening the ability dispersion (and, hence, the dispersion of wages) among educated workers. At the same time, dispersion is narrowed among the less educated.

Our data do indicate an increase in the fraction of the population getting a college degree in Poland over the 1988–1996 period. But whether this mechanism can have a quantitatively important impact on within-group inequality over such a short period is unclear. More promising is the possibility that the large numbers of less educated workers who exited the labor force through early retirement were self-selection on the basis of low ability, and that this selection effect tended to narrow earnings dispersion within low education groups.

The history-dependent explanation for within-group inequality proposed by [Aghion et al. \(2000\)](#) may also help explain this pattern. The idea is that employers use signals such as employment histories, as well as observed attributes like education, to infer worker quality. But employment histories may be less useful in predicting ability in a period of rapid change. The damping effect on within group inequality might be greater for low education workers if it is harder to discriminate among workers on the basis of primary school records than on the basis of cumulative educational records up to and including college.

7. Concluding remarks

This paper has analyzed changes in the structure of earnings in Poland over the period 1985–1996. This period covers the last few years of the pre-transition era, the “big bang” in 1989–1990, and six subsequent years of transition. In our view, it is particularly important to document what happened during the Polish transition, because Poland has been among the most successful of all the transition countries.

As expected, we found that earnings inequality rose substantially during the transition. We also expected to find that most of this increase was attributable to substantial reallocation of labor from the state-owned sector, where wages are compressed because they are not set competitively, to a dynamic private sector that exhibits much greater earnings inequality—a view consistent with the framework in [Aghion and Commander \(1999\)](#), for example. But, to our surprise, we found that the largest single factor driving increased earnings inequality was the growth of *within* sector inequality. Indeed, inequality within the state-owned sector itself grew substantially—by some measures by as much as in the private sector.

The main area where Poland has lagged behind other transition economies is in the pace of privatization of state-owned enterprises (SOEs). Indeed, the state sector still accounted for over 60% of total nonagricultural employment by 1996.³² But a number of authors, including [Pinto et al. \(1993\)](#) and [Commander and Dhar \(1998\)](#), have argued that substantial restructuring of SOEs has occurred even in the absence of privatization. Our findings of increased earnings inequality within the state sector, including substantially increased education premiums, are consistent with this view, since they suggest that the state sector has indeed moved closer to having a competitive wage structure.

Our results indicate that education premiums grew substantially during the transition, by about 50% in the state sector and 100% in the private sector. Presumably, a key factor driving increasing returns to skill in Poland is simply the rising fraction of firms engaged in competitive wage setting. As this fraction grows—both through increase in the size of the private sector and restructuring of SOEs—the relative demand for skilled labor will increase in the economy as a whole. Through this mechanism, education premiums are driven up in both the private and public sectors during the transition.

In contrast to education premiums—which increased to the point where they now appear to be at least as great as those observed in the U.S. and U.K.—experience premiums in Poland have not increased systematically, and remain very low relative to levels observed in the advanced Western economies. This is consistent with the notion of rapid obsolescence of firm- or industry-specific skills during a period of rapid industrial restructuring (see [Svejnar, 1996](#)).

We also find that increases in within-group inequality account for 60% of the overall increase in earnings inequality. But these increases in within-group inequality were much greater for highly educated workers. Models of rising earnings inequality based on technical change that increases returns to both observed and unobserved worker skills typically predict rising within group inequality across all education groups. Thus, it is a

³² The pace of privatization has increased more recently—see [OECD \(2000\)](#).

challenge for models of transition to generate the concentration of growing within group inequality among more educated workers. One potential explanation is based on selection. Large numbers of workers took early retirement during the Polish transition, and these were disproportionately drawn from among workers with low levels of education. If those who retired also tended have low unobserved skills, this would have compressed within group inequality among low education groups.

Finally, we found that the industry wage structure changed dramatically in the early phases of the Polish transition. In 1990–1991, earnings in manufacturing and mining, industries traditionally favored by communist regimes, fell substantially relative to those in higher skilled, more white-collar industries like Finance and Insurance, Public Administration, Science and Technology, Education and Culture, and Health. This led to a massive reallocation of labor out of manufacturing and toward the higher skilled, more white-collar, industries where relative wages rose.

These changes in industry relative wages had important effects on education premiums. In the pre-transition period, better educated workers were relatively concentrated in relatively low wage industries, but this pattern quickly vanished, given the changes in the inter-industry wage structure in 1990–1991. This contributed to rising education premiums (as relative earnings of low education workers in manufacturing fell), and substantially reduced the explanatory power of industry dummies in earnings regressions.

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Appendix A. Accounting for the change in survey frequency after 1992

Most aspects of the HBS, including the two stage sampling scheme and the structure of the survey instrument, were left unchanged during the transition. However, a few refinements were introduced in 1993. While most of these modifications were relatively minor, one important change was that, starting in 1993, households were surveyed for only one month rather than for a full quarter in an attempt to further improve survey response rates. The change in survey frequency has important implications for the purposes of measuring changes in cross-sectional inequality. In this appendix, we develop a technique

for adjusting the 1994–1996 earnings data for the increased variability that may be attributable to the shift from quarterly to monthly reporting.

We begin by assuming the following statistical model for earnings:

$$Y_{ht} = \alpha_{1t} + \beta_t X_{ht} + \sigma_t \epsilon_{ht} \quad (1)$$

where Y_{ht} is labor income of person h in period t , X_{ht} is a vector of individual-specific characteristics used to predict labor income, and ϵ_{ht} is the unpredictable or idiosyncratic component of labor income scaled to have a standard deviation of unity. The time-specific standard deviation that scales the idiosyncratic component of labor income is denoted by σ_t . Our objective is to estimate the increase in σ_t for the 1994–1996 period that is due solely to the switch to a monthly reporting interval.

We begin by estimating Eq. (1) separately for each quarter during 1985–1992 and each month for the period 1994–1996.³³ The variables included in X_{ht} are controls for education level, experience and its square, sex, urban residence, and sector of occupation. A key feature of the results is that the R^2 values drop sharply after the shift to monthly reporting in 1994. Presumably, the bulk of this drop is due to greater idiosyncratic variability of income, as well as greater relative importance of measurement error, when income is reported at monthly rather than quarterly frequencies.

Next, we assume that the standard deviation of the residuals from estimation of Eq. (1) follow the process:

$$\ln \sigma_t = \pi_0 + \pi_1 t + \pi_2 t^2 + \pi_3 t^3 + \pi_4 YB_t + \pi_5 PVTSHR_t + \pi_6 I[t > 96] + \eta_t \\ t = 2, 4, \dots, 95; t = 109, 110, \dots, 144. \quad (2)$$

Here t is a monthly time index. For the years 1985 through 1992, the data are quarterly, so t is assigned as the midpoint of the interval covered by each quarter (that is, $t=2, 4, \dots, 95$). The variable $I[t > 96]$ is an indicator for the post-1992 period in which the data is monthly. Thus, π_6 captures the structural shift in the error standard deviation attributable to the shift to a monthly data frequency. The time polynomials capture the evolution of the error standard deviation over time due to changes in within-group income inequality, controlling for the group characteristics included in X_{ht} . The term YB_t controls for the effect of changes in mean income on the error standard deviation. $PVTSHR_t$ is the share of total employment in the private sector. This share increased markedly after the transition (see Table 1) and, given that earnings inequality tends to be higher in the private sector, could be an important determinant of overall earnings inequality. Finally, η_t captures purely idiosyncratic period-specific changes in income variability. Note that, since we do not have earnings data for 1993, we assume smooth adjustment from the end of 1992 through the beginning of 1994, with no discrete breaks in the intervening year.

After estimating the first and second stage equations, we adjust labor income data for 1994–1996 to account for the increase in the idiosyncratic variance that we estimate

³³ These results are not reported in the paper but are available from the authors. As noted earlier, individual earnings data are not available in the HBS for 1993.

Table A1
Gini coefficients and variance

	All workers	All workers Wage residuals	Private sector	Public sector	Men	Women
1985	0.216	0.181	0.216	0.215	0.206	0.185
1986	0.220	0.185	0.219	0.220	0.209	0.190
1987	0.218	0.182	0.222	0.218	0.206	0.188
1988	0.221	0.188	0.227	0.220	0.213	0.193
1989	0.231	0.202	0.239	0.229	0.222	0.216
1990	0.228	0.200	0.243	0.225	0.222	0.217
1991	0.229	0.194	0.249	0.223	0.227	0.213
1992	0.239	0.207	0.247	0.236	0.235	0.224
1994	0.259	0.214	0.282	0.247	0.261	0.235
1995	0.262	0.217	0.278	0.250	0.261	0.241
1996	0.261	0.214	0.284	0.243	0.261	0.242

	College degree	High school	Vocational training	Primary school
1985	0.191	0.203	0.223	0.212
1986	0.195	0.205	0.229	0.215
1987	0.192	0.207	0.225	0.209
1988	0.211	0.208	0.225	0.207
1989	0.206	0.219	0.234	0.219
1990	0.206	0.212	0.227	0.210
1991	0.209	0.212	0.224	0.203
1992	0.219	0.226	0.236	0.209
1994	0.259	0.238	0.245	0.214
1995	0.256	0.242	0.245	0.233
1996	0.254	0.239	0.241	0.234

Gini coefficients are for real income (levels, not logarithms). Wage residuals are from annual OLS regressions of earnings on a constant, four education dummies, experience and its square, and dummies for gender, urban residence, and employment in the private sector.

occurred solely due to the shift to a monthly reporting period after 1992. We define adjusted income for the 1994–1996 period as:

$$YA_{ht} = \alpha_{1t} + \beta_r X_{ht} + \{\sigma_t / \exp(\pi_6)\} \hat{\epsilon}_{ht} \quad t = 109, 110, \dots, 144. \quad (3)$$

Here $\hat{\epsilon}_{ht}$ is the estimated residual from Eq. (1) and π_6 is our estimate from Eq. (2) of the increase in the log of the residual standard deviation due to the switch to monthly income reporting.

The scale factor $\exp(\pi_6)$ that we estimated was 1.051. This implies that the change in survey frequency results in an estimated increase of about 5% in the residual error variance. Thus, even though employed workers (a majority of whom are salaried) have reasonably stable labor earnings streams, there is still a significant increase in earnings variability in going from a quarterly to a monthly frequency.

Interestingly, the corresponding adjustment factor that we estimated in Keane and Prasad (2002) for total incomes of worker-headed households was 1.179. One would indeed expect the variability of household income (and household labor earnings) to be higher than for individual employed workers. Note that we observe wages for workers only if they are

employed. On the other hand, we do observe zero labor incomes for worker households. This, in addition to variability in employment status among household members, suggests that there is likely to be greater cross-sectional variability in labor earnings and total income of households compared to the sample of employed workers. Furthermore, as the estimated adjustment factors indicate, the increase in cross-sectional variability at higher frequencies is greater for households than for individual employed workers.

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