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Returns to schooling in Russia and Ukraine: A semiparametric approach to cross-country comparative analysis

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Measuring the returns to schooling over an extended period in Russia and Ukraine from 1985 to 2002, we find an increase in both countries but the increase is much bigger in Russia than in Ukraine. To investigate why returns to schooling in Russia and Ukraine diverged over the transition period while the skill composition of employment did not, we compare the Mincerian earnings functions between the two countries and then employ decomposition techniques. Using semiparametric methods, we construct counterfactual wage distributions for university and secondary school graduates in Ukraine using the distributions of Russian characteristics, returns to characteristics, and unobservables. Therefore, we can decompose differences in returns to schooling between the two countries due to differences in the labor market returns, differences in unobservables and differences in labor force composition. We conclude that the price effect is the dominant reason for the observed differences in returns to schooling between these two countries. *Journal of Comparative Economics* **33** (2) (2005) 324–350. University of Michigan, Ann Arbor, MI; IZA, Bonn.

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1. Introduction

This paper contributes to the understanding of the variation in returns to schooling among the countries that have gone through significant economic transformation. Although the returns to schooling have been increasing in the economies that moved from plan to market, we do observe significant variation across countries in the speed of the changes. Russia experienced a sharp increase in returns to schooling within a few years of reform, whereas Ukraine exhibited a very low rate of growth. We investigate the puzzling question of why returns to schooling in Russia and Ukraine diverged so much over the transition period while the skill composition of employment did not. Our study takes advantage of the institutional comparability between these two countries.[1] Russia and Ukraine were part of the Soviet Union until 1991 and shared the same government, institutions, and policies. As we demonstrate, the two countries had remarkably similar wage distributions, earnings structure, educational attainment, labor force composition, and returns to schooling during the pre-reform period. Even now Russia and Ukraine continue to have similar educational systems and workforce characteristics.

Despite this common history and similar initial conditions, the two economies performed quite differently. Ukraine made very few structural reforms until 1997 and only after 1997 did the speed of reforms accelerate and the scope widen (Linn, 2001).[2] In assessing the progress of the transition to a market economy in all twenty-seven countries in the region, the European Bank for Reconstruction and Development consistently scored the results of market reforms less favorably in Ukraine than in Russia (EBRD, 2001). The labor market outcomes of reforms in the two countries are also different.[3] In Russia, overall wage inequality increased sharply, with a significant increase in returns to schooling (Brainerd, 1998; Sabirianova Peter, 2003). In contrast, skill wage inequality in Ukraine did not increase as much over the same period and returns to schooling were among the lowest of the countries of Central and Eastern Europe. This observation of low reform progress

[1] The breakup of Czechoslovakia provides another interesting case for analyzing the divergence in returns to schooling between Czech and Slovak Republics during the reform period (Chase, 1998; Filer et al., 1999).

[2] Typically, Russian economic reforms preceded similar changes in Ukraine. Russia liberalized most domestic prices in January 1992, while Ukraine did so only at the end of 1994. Ukraine introduced a uniform exchange rate two years later after Russia, with full current account convertibility introduced only in 1997. Russia completed mass privatization by July 1994, while Ukraine began its large-scale privatization program at the end of 1994 (EBRD, 2001).

[3] Ukraine lagged behind Russia in labor market reforms as well. Russia abolished the wage grid in the non-public sector in 1991, but Ukraine continued to allocate wages according the old wage grid based on the national agreement between trade unions and the government until 1993. Similarly, Russia abolished a system of penalties on the growth of wage fund in 1995 while Ukraine did so only at the end of 1996. Until 2004, Ukraine had consistently higher marginal personal income tax rates and indirect income taxes paid by enterprises than Russia.

and concurrent low returns to schooling in Ukraine is consistent with the recent finding of Fleisher et al. (2005) on the positive effect of the speed of market reforms on a country's returns to schooling.

Our approach in analyzing the sources of cross-country differences in returns to schooling is to compare the Mincerian earnings functions between the two countries and then to use semiparametric decomposition techniques following Juhn et al. (1993). We construct counterfactual distributions of log wages for university and secondary school graduates in Ukraine using the distributions of Russian characteristics, returns to characteristics, and unobservables. These counterfactual distributions provide an estimate of the distributions of Ukrainian log wages that would have prevailed if Ukraine had the same features as Russia. This allows us to decompose differences in returns to schooling between the two countries into shares due to differences in the labor market returns, i.e., the price effect, differences in unobservables, i.e., the residual effect, and differences in the labor force composition, i.e., the composition effect.

In our comparative analysis, we use Russian and Ukrainian Longitudinal Monitoring Surveys covering a period from 1985 to 2002. In addition to the institutional comparability between the two countries, our study also benefits from the definitional comparability between the two surveys. Most of the key variables have similar definitions and the estimated earnings functions have the same specifications. Thus, we avoid a common problem in cross-country studies, in which the differences in the estimated parameters are influenced by discrepancies in the quality of data, estimation methods, and definitions of variables.[4] In the next section, we describe in detail the data and variables used in the empirical analysis. Section 3 provides the comparative analysis of conventional earnings functions and discusses the robustness of estimated returns to schooling with regard to the choice of specifications, variables, and methods used. Section 4 uses semiparametric methods to decompose the sources of cross-country differences in returns to schooling. Section 5 concludes with a summary of the findings.

2. Russian and Ukrainian Longitudinal Monitoring Surveys

The data are pooled from two household surveys, namely Russian Longitudinal Monitoring Survey (RLMS) and Ukrainian Longitudinal Monitoring Survey (ULMS).[5] We use the second wave of RLMS that started in 1994; it selected 4781 dwelling units by a three-stage stratified clustered sampling method and 3973 households responded. In the subsequent years of survey, i.e., 1995–1996, 1998, and 2000–2002, the new households that moved to the initially sampled dwellings were added and the old households that

[4] Behrman and Rosenzweig (1994) and Srinivasan (1994) provide a useful discussion of this problem in emerging markets.

[5] RLMS was organized by Barry Popkin and conducted by the Consortium led by the Carolina Population Center in collaboration with the Institute of Sociology at the Russian Academy of Sciences, Moscow. ULMS was organized by Hartmut Lehmann and carried out by the Consortium led by the Institute for the Study of Labor (IZA), Bonn in collaboration with the Kyiv International Institute of Sociology. We are thankful to both teams for their excellent work.

moved from the original sample to new addresses were included, whenever possible. The sample size varies from year to year; individuals who completed the adult questionnaire numbered 8893 in 1994, 8342 in 1996, 8701 in 1998, 9074 in 2000, and 10,497 in 2002. To ensure the representativeness of cross-sections, we exclude those respondents who moved from the original sample.[6] Our sample consists of 8122 respondents in 1996, 7894 in 1998, 7568 in 2000, and 7875 in 2002 so that we base our wage analysis on a sample of 3384 to 4415 employed adults for whom we have complete information on wages, education, and demographic characteristics. The RLMS 2000 also contains a series of retrospective questions regarding jobs held in 1985 and 1990. The number of persons who responded to the questions on wages was 4230 in 1985 and 3976 in 1990.

The Ukrainian survey is based on a national stratified random sample of 4096 households. Unlike RLMS, ULMS started only in 2003. To include labor force data in Ukraine in the 1990s, ULMS gathered employment histories for 1986, 1991, and continuously from 1997 to 2003. The response rate was 66% for households and 87% for individuals within the households, resulting in 8641 individuals of age 15 to 72 years participating in the survey. The sample of employed persons with non-missing values on wages, education, and demographic characteristics ranges from 2958 in 1998 to 4197 in 1986.

Both surveys contain rich information on household and individual characteristics, although the focus is somewhat different in each. RLMS focuses more on household behavior and has extensive sections on household income and expenditures, health, nutrition, children and women issues, whereas ULMS devotes a significant portion of its questionnaire to the retrospective histories of employment, education, and migration. Despite these differences, the two surveys provide a consistent set of the individual characteristics, including individual earnings, hours of work, education, demographics, job tenure, and characteristics of the primary employer, e.g., ownership and size.

The definitions of all variables used in the empirical analysis are provided in Appendix Table A.1. Most of the individual attributes have identical definitions in both surveys, e.g. gender, age, potential labor market experience, job tenure, average weekly hours of work, and capital. In a few cases, an original variable has been modified to make it comparable across the surveys. For example, the continuous variable of the employer size from RLMS has been recoded into a categorical variable having the same size categories as in ULMS. Detailed information on firm ownership from ULMS has been aggregated into the three broad categories of ownership available in RLMS, namely foreign (including domestic firms with some foreign capital), private (including self-employed, cooperatives, fully and partially privatized enterprises, and newly established private enterprises), and state (including budgetary organizations, state enterprises, local municipal enterprises, and state and collective farms).

Both Russian and Ukrainian data contain detailed information on formal schooling, including the type of schools, actual years of studies, degrees obtained, and the date of school completion. At least two alternative measures of the years of schooling can be

[6] The estimates of the returns to schooling are not affected by excluding respondents who moved from the original sample. We re-estimated the standard Mincerian earnings function on a larger sample by adding a dummy for movers and an interaction term with years of schooling. In all years, we find no statistically significant, at the 10% level, differences in wages and returns to schooling due to respondents' moving from the original sample.

constructed for the years of survey, namely, the actual years of studies from all schools attended and the adjusted years of schooling that are imputed from the number of years required for the highest degree obtained. However, for the retrospective years 1985 to 1991 only the latter measure can be imputed accurately using the date of school completion so that we employ this second measure in most of our empirical analysis. The correlation between these two measures of schooling is relatively high; in 2002, the simple correlation coefficient is 0.900 in Russia and 0.895 in Ukraine.

The dependent variable is the log of monthly contractual (accrued) wages after taxes at the primary job. Using the contractual wage in the earnings functions is preferred over wages received in some short reference period, e.g., a month, especially during periods of mass wage delays and high volatility in wage payments, which both countries experienced in the 1990s.[7] Wages actually received in the last month are often zero as the wage debt accumulates; it can become much higher than the contractual wage when the debt is paid back. In ULMS, the measure of net contractual wages is available for all years for both employees and self-employed. Wages received in a different currency are converted into Ukrainian hryvnyas (UAH).[8] In RLMS, this measure is available only for employees and only from 1998 to 2002. For 1994 to 1996, we follow the method of Earle and Sabirianova Peter (2002) and impute the contractual wage for workers with wage arrears as the ratio of the total wage debt to the number of monthly wages owed. For workers without wage arrears, the contractual wage is considered to be the actual after-tax monthly wage received in cash or in kind in the last 30 days from the primary job. This measure of actually paid earnings is also used for self-employed from 1994 to 2002.

For the Soviet period, the definitions of wages in both countries are the same so that the only main concern is the possibility of recall bias. Although people may not remember the wages they received 17 or even 10 years ago, wage is the dependent variable so that recall bias should not affect the results as long as it can be assumed to be an additive white noise. In addition, the Soviet practice of wage payments according to the rigid wage grid, nearly zero inflation and strong attachment of a Soviet worker to one job are likely to reduce recall error (Münich et al., 2005a). Importantly, three of the four years selected are memorable, pivotal points in the Soviet history. In 1985, Gorbachev came to power and perestroika began, in 1986, the Chernobyl catastrophe in Ukraine shook the world, and, in 1991, Gorbachev resigned, the Soviet Union ended, and Russia and Ukraine began their new independent history. Figure 1 indicates that the shape of wage distributions in 1985/1986 is similar for the two countries, which is remarkable given that the two samples are drawn independently in different years and that the recall period is one year and half longer in Ukraine than in Russia. The mean wages from the surveys, denoted w_{SAM}, are close to the mean wages from the national statistical yearbooks, denoted w_{NSY}, for corresponding years; $w_{\mathrm{SAM},85} = 207$ and $w_{\mathrm{NSY},85} = 199$ for Russia and $w_{\mathrm{SAM},86} = 173$

[7] The contractual wage also has shortcomings since it implicitly assumes that costs associated with the delay of payments are zero and that all arrears will be paid back. Ideally, we would like to use wages actually received during a longer time period, e.g., 6 or 12 months. The ULMS data set provides such a variable but only in the last year.

[8] Ukraine used rubles until 1992, karbovancy from 1992 to 1996, and hryvnyas after 1996. People may also receive wages in a foreign currency, typically US dollars.

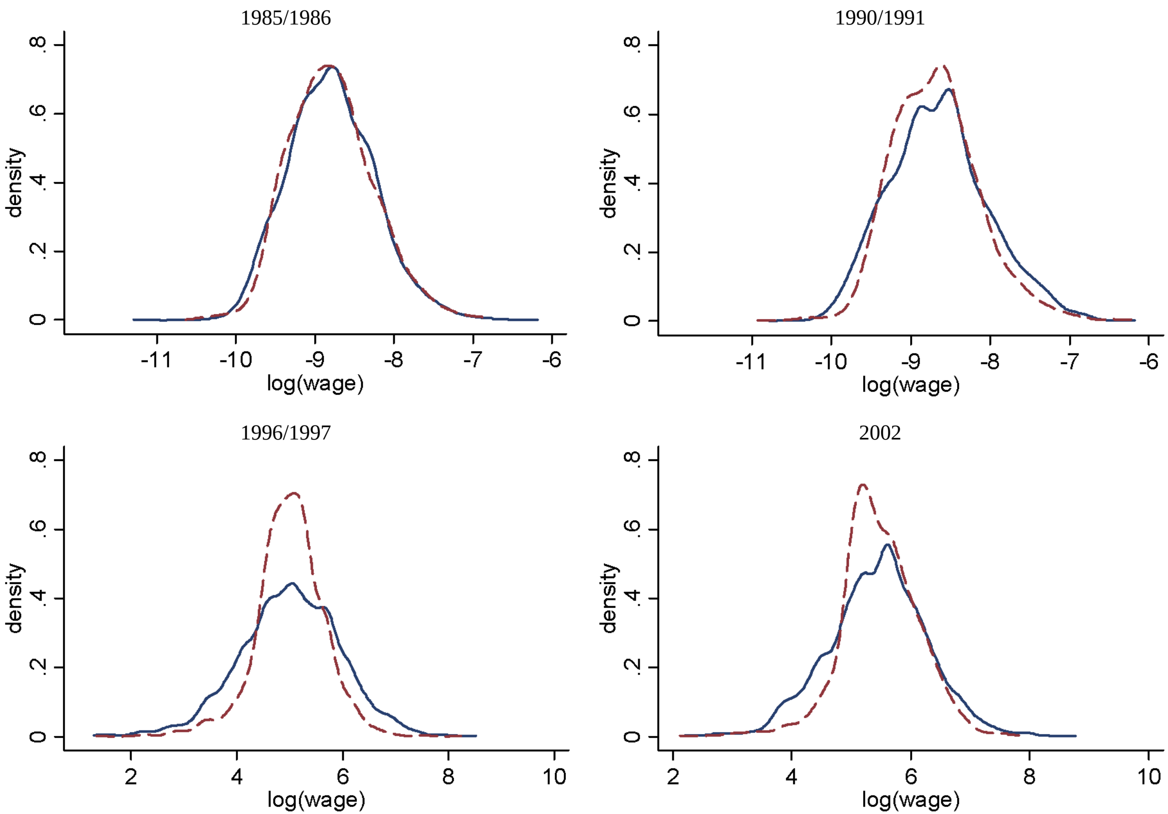

Fig. 1. Wage distributions, Russia and Ukraine. *Notes:* In all graphs, broken and solid lines correspond to Ukraine and Russia, respectively. Wages are rescaled so that the mean log wage in Russia is equal to the mean log wage in Ukraine for each year. The plotted densities use the Parzen kernel with bandwidth of 0.45.

and $w_{\text{NSY},86} = 174$ for Ukraine. Hence, we find it plausible to approximate the true wage distributions in the Soviet period by the wages reported retrospectively. Moreover, even if errors of recall are present in the data, they should not bias the results in the direction of one country or the other.[9]

Because older age groups are under-represented in the retrospective surveys, we restrict our analysis to the prime age group 15 to 59 to reduce the potential effect of the mortality-related sample attrition on the estimates of returns to schooling. For the Soviet period, we use the sample weights based on the 1989 USSR Census, with under-represented groups receiving larger weights. We also use the sample weights for more recent years but the results are not statistically different.

Table 1 and Appendix Table A.2 report the summary statistics for both surveys. Russia and Ukraine exhibit similar labor force characteristics during the pre-reform and reform periods. Mean adjusted years of schooling are identical in Russia and Ukraine in 1985/1986 and differ only by 0.6 years in 2002.[10] The average length of the workweek is between 41 and 43 hours and the mean labor market experience is around 21 years in both countries. However, tenure is longer in Ukraine than in Russia by 0.8 to 1.6 years. Although certain differences are noticeable with respect to firm characteristics, e.g., Ukraine has a higher share of the employed in very large enterprises and in the state sector, an increase in the share of workers in private and foreign-owned firms and small businesses is apparent in both countries. Appendix Table A.2 suggests that supply changes are unlikely to explain the different time paths in returns to schooling between the two countries because of similar dynamics in workers' educational attainment. In both countries, the share of workers with university degrees has been growing continuously at the same rate and the share of low-education workers is declining. Finally, Fig. 1 documents that starting from essentially identical shapes in 1985/1986, Russian and Ukrainian wage distributions diverged significantly over the transition period, with Russia having considerably higher levels of overall wage inequality.

3. Earnings function analysis

In this section, we compare the estimates of the Mincerian earnings functions in Russia and Ukraine and present the results of sensitivity analysis to establish the robustness of the estimated returns to schooling with regard to the choice of specifications, variables, and methods used. We begin by estimating the basic Mincerian earnings function with a

[9] Determining the sign of the bias generated by the recall errors in variables is difficult because the recall error is likely to be mean reverting, especially in wages (see Kim and Solon, 2005). We expect, however, the size of the recall error to be negatively correlated with memory and so that the error should be increasing with age and decreasing with education. Hence, the coefficients on schooling are likely to be less downwardly biased than coefficients on tenure and labor market experience. In any case, we have no reason to expect the bias to be different in the two countries.

[10] The reported differences in the actual years of schooling between the two countries could be due to the fact that ULMS specifically asks respondents not to count interruptions during the study, such as maternity leave or required army service, while RLMS respondents might have included these breaks in the total years of schooling.

Table 1
Descriptive statistics

Panel A: Russia

	1985		1990		1996		1998		2000		2002	
	Mean	St. dev.	Mean	St. dev.	Mean	St. dev.	Mean	St. dev.	Mean	St. dev.	Mean	St. dev.
Log(wages)	−1.697	0.533	−1.367	0.628	6.173	0.951	6.462	0.845	7.115	0.908	7.856	0.839
Schooling (adjusted years)	10.742	3.070	11.212	2.748	11.774	2.423	11.879	2.342	11.911	2.236	12.417	2.329
Schooling (actual years)					12.317	2.770	12.499	2.694	12.522	2.574	12.658	2.716
Female	0.482	0.500	0.475	0.499	0.505	0.500	0.501	0.500	0.491	0.500	0.503	0.500
Tenure (years)					7.933	8.691	7.680	8.517	7.185	8.207	7.627	9.068
Experience (years)	21.390	11.984	20.901	11.625	20.407	10.856	20.431	10.703	20.562	10.500	20.997	11.654
Log(hours)					3.680	0.405	3.702	0.327	3.750	0.313	3.749	0.311
Hours missing					0.125	0.331	0.016	0.125	0.026	0.158	0.013	0.115
Capital city	0.033	0.180	0.031	0.174	0.068	0.252	0.051	0.220	0.033	0.180	0.066	0.249
Ownership												
Private					0.278	0.448	0.295	0.456	0.337	0.473	0.410	0.492
Foreign					0.032	0.175	0.033	0.179	0.039	0.194	0.043	0.202
State					0.541	0.498	0.517	0.500	0.486	0.500	0.467	0.499
No information					0.149	0.356	0.155	0.361	0.138	0.345	0.080	0.272
Employer size (no. of persons)												
1–10					0.108	0.310	0.112	0.316	0.141	0.348	0.163	0.369
10–50					0.199	0.399	0.194	0.396	0.183	0.387	0.208	0.406
50–100					0.099	0.299	0.107	0.309	0.097	0.296	0.093	0.290
100–500					0.166	0.373	0.174	0.379	0.170	0.376	0.174	0.379
500–1000					0.043	0.202	0.057	0.232	0.061	0.239	0.063	0.244
>1000					0.083	0.275	0.078	0.268	0.091	0.288	0.089	0.285
Size is missing					0.302	0.459	0.278	0.448	0.257	0.437	0.209	0.407
Sample size, *N*	4111		3776		3497		3332		3169		3341	

(*continued on next page*)

Table 1 (*continued*)

Panel B: Ukraine

	1986		1991		1997		1998		2000		2002	
	Mean	St. dev.	Mean	St. dev.	Mean	St. dev.	Mean	St. dev.	Mean	St. dev.	Mean	St. dev.
Log(wages)	−8.791	0.528	−8.666	0.626	5.010	0.634	5.106	0.623	5.312	0.607	5.481	0.640
Schooling (adjusted years)	10.750	2.892	11.157	2.559	11.612	2.361	11.688	2.275	11.822	2.239	11.849	2.217
Schooling (actual years)					10.915	2.077	10.977	2.011	11.074	2.012	11.118	2.028
Female	0.496	0.500	0.483	0.500	0.537	0.499	0.539	0.499	0.538	0.499	0.530	0.499
Tenure (years)	10.406	9.931	10.306	9.984	9.248	9.617	9.379	9.730	8.770	9.466	8.464	9.326
Experience (years)	21.241	11.860	21.331	11.797	21.088	11.104	21.183	10.793	20.884	10.607	21.198	10.706
Log(hours)											3.659	0.324
Hours missing											0.097	0.296
Capital city	0.037	0.190	0.036	0.187	0.058	0.235	0.058	0.233	0.065	0.246	0.063	0.243
Ownership												
Private	0.104	0.305	0.154	0.361	0.224	0.417	0.242	0.428	0.276	0.447	0.298	0.457
Foreign	0.005	0.071	0.009	0.095	0.015	0.120	0.013	0.112	0.015	0.120	0.016	0.127
State	0.885	0.320	0.828	0.377	0.744	0.436	0.723	0.448	0.683	0.465	0.652	0.476
No information	0.007	0.082	0.009	0.093	0.018	0.132	0.022	0.147	0.026	0.159	0.033	0.180
Employer size (no. of persons)												
1–10	0.074	0.261	0.089	0.285	0.106	0.308	0.104	0.305	0.122	0.327	0.148	0.355
10–50	0.092	0.290	0.103	0.304	0.143	0.350	0.159	0.366	0.176	0.381	0.208	0.406
50–100	0.055	0.229	0.060	0.237	0.085	0.279	0.094	0.292	0.103	0.303	0.113	0.317
100–500	0.090	0.287	0.112	0.315	0.143	0.350	0.152	0.359	0.168	0.374	0.187	0.390
500–1000	0.031	0.174	0.031	0.172	0.042	0.201	0.044	0.205	0.051	0.220	0.062	0.241
>1000	0.073	0.260	0.089	0.284	0.117	0.321	0.126	0.332	0.136	0.343	0.150	0.357
Size is missing	0.584	0.493	0.517	0.500	0.365	0.481	0.321	0.467	0.245	0.430	0.133	0.339
Sample size, N	4191		3528		2946		2812		2925		3289	

Note. The sample for each country and year consists of observations with non-missing values for the variables used in the basic Mincerian wage function.

standard set of covariates available for both countries and all years specified as[11]:

$$\ln w_{it} = \beta_0 + \beta_1 sch_{it} + \beta_2 exp_{it} + \beta_3 exp_{it}^2 + \beta_4 female_{it} + \beta_5 capital_{it} + \varepsilon_{it}, \qquad (1)$$

where i indexes individuals, t indexes time, w_{it} is monthly contractual wages after taxes at the primary job, sch_{it} is adjusted years of schooling, exp_{it} is years of potential labor market experience, $female_{it}$ is a dummy variable indicating if an individual is female, $capital_{it}$ is a dummy variable indicating if an individual i lives in the capital city, and ε_{it} is an independently distributed error term.

The Ordinary Least Squares (OLS) estimates of Eq. (1) for Russia and Ukraine are presented in Table 2. Until 1991, returns to schooling are similar for Russia and Ukraine at

Table 2
Basic Mincerian earnings functions, OLS

Panel A: Russia

	1985	1990	1996	1998	2000	2002
Schooling (adjusted years)	0.028***	0.039***	0.081***	0.091***	0.093***	0.092***
	(0.003)	(0.004)	(0.007)	(0.006)	(0.007)	(0.006)
Experience (years)	0.019***	0.027***	0.014***	0.029***	0.045***	0.030***
	(0.003)	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)
Experience^2/1000	−0.378***	−0.554***	−0.335***	−0.629***	−0.991***	−0.724***
	(0.055)	(0.072)	(0.119)	(0.108)	(0.129)	(0.119)
Female	−0.424***	−0.401***	−0.473***	−0.530***	−0.520***	−0.473***
	(0.016)	(0.020)	(0.030)	(0.027)	(0.030)	(0.026)
Capital	0.011	0.095	0.614***	0.537***	0.634***	0.630***
	(0.044)	(0.058)	(0.060)	(0.061)	(0.082)	(0.053)
N	4111	3776	3497	3332	3169	3341
R^2	0.19	0.15	0.13	0.17	0.15	0.18

Panel B: Ukraine

	1986	1991	1997	1998	2000	2002
Schooling (adjusted years)	0.034***	0.039***	0.037***	0.039***	0.037***	0.045***
	(0.004)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Experience (years)	0.012***	0.016***	0.019***	0.016***	0.016***	0.019***
	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Experience^2/1000	−0.230***	−0.320***	−0.436***	−0.395***	−0.415***	−0.502***
	(0.063)	(0.078)	(0.084)	(0.088)	(0.090)	(0.092)
Female	−0.418***	−0.434***	−0.423***	−0.423***	−0.413***	−0.398***
	(0.016)	(0.021)	(0.022)	(0.022)	(0.021)	(0.021)
Capital	0.159***	0.167***	0.267***	0.274***	0.301***	0.285***
	(0.044)	(0.057)	(0.046)	(0.047)	(0.042)	(0.043)
N	4191	3528	2946	2812	2925	3289
R^2	0.20	0.15	0.14	0.14	0.14	0.13

Notes: (i) Dependent variable is log of monthly contractual wages after taxes at the primary job. (ii) Robust standard errors are in parentheses. (iii) The sample weights are applied in Russia for all years and in Ukraine for 1986 and 1991. (iv) Constant term is estimated but not reported.

*** Significance at the 1% level.

[11] We restrict our analysis to the years that are available for both countries, namely 1985 (1986), 1990 (1991), 1996 (1997), 1998, 2000, and 2002.

2.8 to 3.4% in 1985/1986[12] and 3.9% in 1990/1991. With the demise of the Soviet Union, schooling returns began to diverge significantly; in Russia, returns increased sharply to 8.1% in 1996 and then to 9.2% in 2002, while in Ukraine they barely changed reaching only 4.5% in 2002. As in a typical Mincerian earnings equation, the estimated returns to potential experience are concave. However, compared to the estimates for the US, the wage-experience profiles are flatter and average returns to labor market experience are relatively small.[13] The small experience effect may be due to by the changing nature of the transition economies in which younger, more mobile, and more adaptive people are rewarded. The experience profile is less concave in Ukraine than in Russia.

The male wage premium is significantly larger in Russia than in Ukraine. Although the gender wage gap fell in Russia from 53% in 1998 to 47% in 2002, this difference is still much higher than it was during the Soviet period in both Russia and Ukraine and during the transition period in Ukraine (40 to 43%). The wage premium for living in a capital city is high in both countries. However, this premium increases from 16% in 1986 to 27% in 1997 and remains approximately constant between 27 and 30% from 1997 to 2002 in Ukraine, but the premium for living in Moscow exhibits a sharp increase from 10% in 1990 to 61% in 1996 and remains above 60% from 2000 onwards.[14]

As a complement to OLS estimates, we present the estimates of returns to schooling obtained from the series of quantile regressions following Koenker and Bassett (1978). While the OLS method produces only mean prices of observable characteristics, quantile regressions can produce the whole distribution of returns to schooling. Hence, we investigate whether observed changes in schooling returns are uniform or concentrated in certain groups. Formally, we estimate the following basic Mincerian function:

$$Q_k(\ln w_{it}|X_{it}) = \beta_0^{(k)} + \beta_1^{(k)} sch_{it} + \beta_2^{(k)} exp_{it} + \beta_3^{(k)} exp_{it}^2 + \beta_4^{(k)} female_{it} + \beta_5^{(k)} capital_{it}, \tag{2}$$

where $Q_k(\ln w_{it}|X_{it})$ denotes the kth percentile of distribution of log wages conditional on the covariate matrix X_{it} and $\beta_j^{(k)}$ is the kth percentile estimate of the slope of variable j. For each percentile k, country, and period, we estimate Eq. (2) and plot the obtained distributions of returns to schooling in Fig. 2. The growing differences in returns to schooling between Russia and Ukraine are apparent. The cross-country differences are more pronounced in the middle of distribution in 1996/1997. By 2002 the bottom percentiles in Russia exhibit the largest increase in returns to schooling implying that having additional education at the bottom of wage distribution improves people's welfare significantly.

[12] The difference in returns to schooling between the two countries in 1985/1986 is not statistically different from zero.

[13] This result is consistent with earlier studies by Flanagan (1998) and Rutkowski (1997), who also document low returns to labor market experience in Czech Republic and Poland.

[14] The relatively low Moscow premium during the Soviet period could be explained by the fact that, unlike other Soviet republics, Russia had many territories in which workers were compensated for living in unfavorable climate conditions. The base Moscow salary was simply multiplied by the regional wage coefficient, which compressed the average Moscow premium. The high premium afterwards may result from the system of living permits that drives up overall wages in the capital.

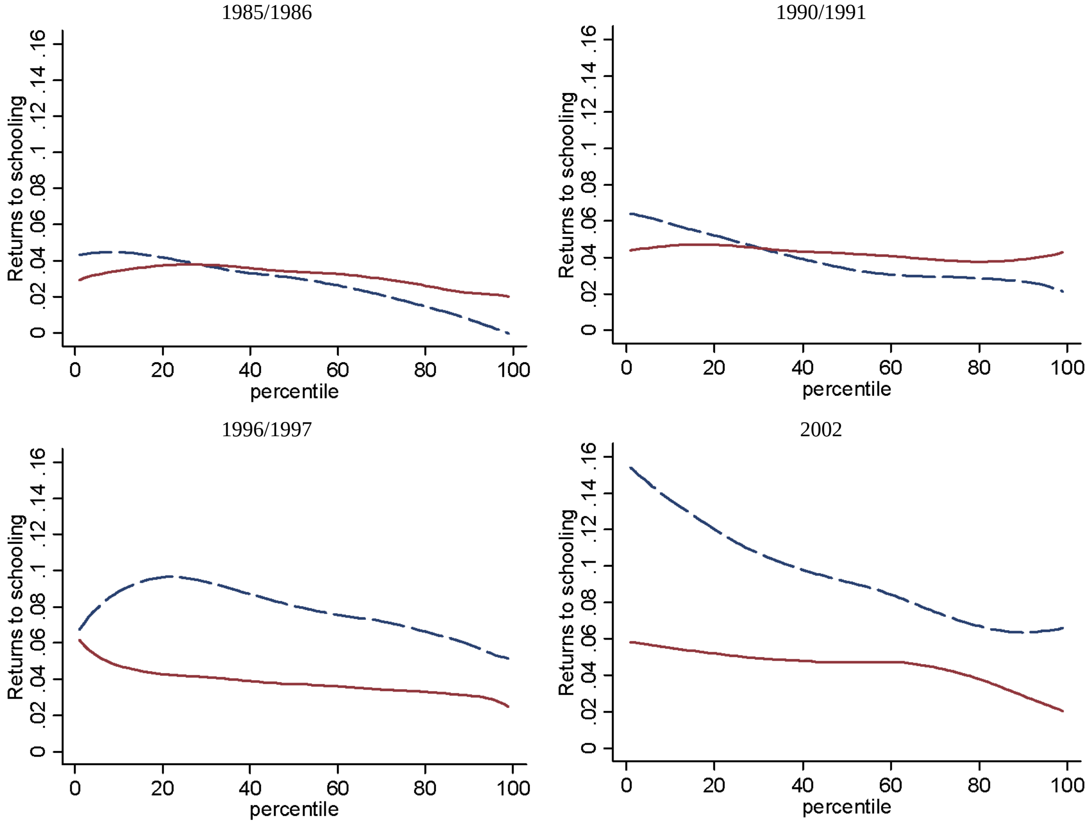

Fig. 2. Distribution of the returns to schooling, Russia and Ukraine, quantile estimates. *Notes:* In all graphs, broken and solid lines correspond to Ukraine and Russia, respectively. The estimated equation is the basic Mincerian specification shown in Table 2. A locally weighted regression (lowess) with bandwidth 0.5 is used to smooth percentile estimates in the figure.

Based on this first analysis of the data, we conclude that Russia and Ukraine had very similar shapes of wage distributions, composition of labor force, returns to schooling, and returns to other observable characteristics during the Soviet period. However, within a few years of reforms the differences in prices of observable characteristics, including returns to schooling, became apparent. Below we investigate further this divergence using a larger set of explanatory variables that are available from 1996 to 2002. In particular, we estimate the following specification:

$$\ln w_{it} = \beta_0 + \beta_1 sch_{it} + \beta_2 exp_{it} + \beta_3 exp_{it}^2 + \beta_4 female_{it} + \beta_5 capital_{it} + \beta_6 ten_{it} + \beta_7 ten_{it}^2 + \sum_{n=1}^{q} \alpha_k own_{n,it} + \sum_{m=1}^{p} \gamma_m size_{m,it} + \varepsilon_{it}, \tag{3}$$

where ten_{it} is years of tenure at the primary job, $own_{n,it}$ is a set of dummies for state, private, and mixed ownership types, $size_{m,it}$ is a set of dummies for the (employment) size categories by employment of the firm individual i works.[15]

The estimates of the augmented Mincerian earnings function are presented in Table 3. The overall trend and the levels of returns to schooling remain qualitatively the same as in the basic Mincerian earnings function. The rate of return increases sharply in Russia in 1998 but hardly changed in Ukraine until 2002. Although controlling for tenure and firm characteristics does not affect returns to labor market experience, it reduces gender wage differences and the premium for living in a capital city by around five percentage points in both countries. The insignificant tenure effect might have been expected, especially during the early reform period, as accumulated firm-specific human capital becomes obsolete in the new economic environment. Remarkably, after ten years of transition the estimated return to tenure is only 0.5% in Russia and 0.8% in Ukraine. Perhaps workers with long tenures continue to be associated with inefficient state firms and lack up-to-date skills.

Firm characteristics contribute significantly to explaining variation in wages, with important differences across ownership types. In both countries, foreign-owned firms pay the highest wages ceteris paribus, followed by private firms, while the state sector has the lowest wages. However, the non-state/state wage gap is somewhat larger in Russia than in Ukraine. Specifically, workers in foreign-owned enterprises earn 42 to 54% more in Russia but only 39 to 45% more in Ukraine relative to state-owned enterprises (SOEs). Likewise, private-owned firms pay their workers 30 to 40% and 16 to 26% more in Russia and Ukraine, respectively, compared to SOEs. Our estimates also demonstrate a significant employer size effect on wage.[16]

[15] We also estimated more flexible functional forms for the earnings allowing the returns to schooling to vary by gender and by ownership type. In both countries and in all years, we can not reject the null hypothesis that the coefficients on interactions of years of schooling with dummies for ownership type and gender equal zero at standard significance levels. The lack of variation in returns to schooling by ownership type is consistent with the hypothesis that an increase in schooling returns in transition economies is common for all sectors and is driven by the market rather than by ownership effects (Münich et al., 2005b).

[16] This result is consistent with the positive size-wage gap documented in the US literature. Traditional explanations of the positive size-wage gap appeal to higher productivity of large firms, selection of better workers, higher monitoring costs, rent-sharing, and efficiency wages to prevent shirking (see Oi and Idson, 1999).

Table 3
Augmented Mincerian earnings functions, OLS

	Russia				Ukraine			
	1996	1998	2000	2002	1997	1998	2000	2002
Schooling	0.079***	0.091***	0.094***	0.097***	0.036***	0.040***	0.038***	0.048***
(adjusted years)	(0.007)	(0.006)	(0.007)	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)
Experience (years)	0.013**	0.025***	0.042***	0.027***	0.021***	0.016***	0.015***	0.018***
	(0.006)	(0.005)	(0.006)	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)
Experience^2/1000	−0.290**	−0.553***	−0.881***	−0.639***	−0.438***	−0.362***	−0.367***	−0.456***
	(0.126)	(0.111)	(0.131)	(0.123)	(0.089)	(0.093)	(0.094)	(0.094)
Female	−0.448***	−0.486***	−0.448***	−0.426***	−0.386***	−0.383***	−0.377***	−0.347***
	(0.031)	(0.027)	(0.030)	(0.026)	(0.022)	(0.022)	(0.021)	(0.021)
Capital	0.552***	0.490***	0.602***	0.573***	0.226***	0.240***	0.272***	0.231***
	(0.059)	(0.059)	(0.080)	(0.052)	(0.045)	(0.045)	(0.041)	(0.041)
Tenure (years)	0.003	0.007	0.001	0.005	−0.005	−0.001	0.005	0.008**
	(0.006)	(0.005)	(0.006)	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)
Tenure^2/1000	−0.068	−0.206	−0.111	−0.160	0.127	0.086	−0.096	−0.153
	(0.191)	(0.169)	(0.194)	(0.166)	(0.112)	(0.112)	(0.116)	(0.119)
Ownership								
Private	0.299***	0.364***	0.401***	0.325***	0.161***	0.219***	0.216***	0.262***
	(0.036)	(0.031)	(0.035)	(0.029)	(0.027)	(0.027)	(0.026)	(0.026)
Foreign	0.486***	0.461***	0.423***	0.541***	0.389***	0.408***	0.424***	0.446***
	(0.086)	(0.076)	(0.077)	(0.065)	(0.093)	(0.096)	(0.086)	(0.081)
Employer size (no. of persons)								
10–50	−0.048	−0.046	−0.100*	−0.067	−0.009	−0.028	−0.032	0.046
	(0.058)	(0.049)	(0.053)	(0.043)	(0.043)	(0.043)	(0.038)	(0.035)
50–100	−0.046	0.060	0.031	0.095*	−0.013	0.014	0.032	0.114***
	(0.067)	(0.057)	(0.062)	(0.054)	(0.049)	(0.049)	(0.044)	(0.041)
100–500	0.036	0.074	0.091*	0.090**	0.098**	0.108**	0.116***	0.220***
	(0.060)	(0.051)	(0.055)	(0.046)	(0.044)	(0.043)	(0.040)	(0.037)
500–1000	0.229***	0.290***	0.155**	0.211***	0.149**	0.153**	0.188***	0.283***
	(0.086)	(0.068)	(0.072)	(0.061)	(0.062)	(0.061)	(0.054)	(0.050)
>1000	0.054	0.216***	0.286***	0.236***	0.245***	0.291***	0.314***	0.450***
	(0.071)	(0.062)	(0.064)	(0.055)	(0.046)	(0.046)	(0.042)	(0.039)
N	3413	3275	3139	3284	2932	2802	2916	3278
R^2	0.15	0.22	0.21	0.23	0.17	0.19	0.20	0.20

Notes: (i) Dependent variable is log of monthly contractual wages after taxes at the primary job. (ii) Robust standard errors are in parentheses. (iii) Sample weights are applied in Russia. (iv) The sample is restricted to ages 15 to 59. (v) The omitted categories are 1–10 for employer size and state for ownership. (vi) The intercept and two dummy variables for missing employer size and missing ownership are included but their coefficients are not reported.

* Significance at the 10% level.
** Idem., 5%.
*** Idem., 1%.

To check the sensitivity of these estimates of returns to schooling, we relax sample restrictions, employ different definitions of the key variables, and include other controls to the baseline equation. Table 4 presents the results of the sensitivity analysis. The estimates are robust to the sample weights in both countries and to the age restrictions of the sample in ULMS, which has 72 years as an upper bound. When older age groups are included in

Table 4
Sensitivity analysis of the estimated returns to schooling

Panel A: Russia

	1985	1990	1996	1998	2000	2002
Without survey weights	0.027	0.039	0.081	0.094	0.097	0.096
	(0.003)	(0.004)	(0.007)	(0.006)	(0.007)	(0.006)
Without restrictions on age	0.027	0.037	0.077	0.083	0.086	0.089
	(0.003)	(0.004)	(0.006)	(0.006)	(0.006)	(0.006)
	[4220]	[3964]	[3676]	[3537]	[3374]	[3531]
Schooling (actual years)			0.068	0.078	0.080	0.080
			(0.006)	(0.005)	(0.006)	(0.005)
			[3469]	[3298]	[3159]	[3310]
With wages actually received last month			0.065	0.090	0.092	0.086
			(0.008)	(0.009)	(0.008)	(0.007)
			[2445]	[2326]	[2649]	[2906]
With log of hourly wage rate			0.085	0.097	0.102	0.101
			(0.007)	(0.006)	(0.007)	(0.006)
			[3061]	[3281]	[3090]	[3297]
With industry dummies	0.032	0.042	0.079	0.091	0.093	
	(0.003)	(0.004)	(0.006)	(0.006)	(0.007)	
	[3982]	[3679]	[3477]	[3316]	[3132]	
With district fixed effects	0.023	0.030	0.063	0.069	0.073	0.074
	(0.003)	(0.004)	(0.006)	(0.005)	(0.006)	(0.005)

Panel B: Ukraine

	1986	1991	1997	1998	2000	2002
Without survey weights	0.031	0.039	0.037	0.039	0.036	0.046
	(0.003)	(0.004)	(0.005)	(0.005)	(0.005)	(0.005)
Without restrictions on age	0.034	0.039	0.040	0.040	0.035	0.046
	(0.004)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
	[4192]	[3564]	[3073]	[2945]	[3099]	[3494]
Schooling (actual years)			0.041	0.046	0.047	0.055
			(0.005)	(0.006)	(0.005)	(0.005)
			[2914]	[2779]	[2890]	[3245]
With wages actually received last month						0.049
						(0.005)
						[3066]
With log of hourly wage rate						0.052
						(0.005)
						[2968]
With wages actually received over the last six months						0.048
						(0.005)
						[2570]
Based on IV estimation	0.072	0.092	0.088	0.103	0.112	0.121
	(0.010)	(0.014)	(0.014)	(0.014)	(0.014)	(0.013)
	[3764]	[3193]	[2722]	[2604]	[2731]	[3058]
With parents' background	0.031	0.034	0.031	0.031	0.028	0.038
	(0.004)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
	[3766]	[3196]	[2736]	[2615]	[2746]	[3061]

Notes: (i) Robust standard errors are in parentheses. (ii) All coefficients are significant at 1%. (iii) The number of observations is in brackets if it is different from Table 2.

RLMS, the estimated rates of return decline by 0.1 to 0.8 percentage points. Alternative definitions of schooling and wages do change the schooling returns but do not affect the overall trends and conclusions. Using actual years of schooling instead of adjusted years of schooling raises the estimates of returns to schooling by 0.4 to 1.0 percentage points for Ukraine and decreases the estimates by 1.2 to 1.3 percentage points for Russia. The wage measure that we criticized earlier for its non-random volatility, namely actually received last month, reduces the baseline estimates by 0.1 to 1.6 percentage points in Russia and increases it by 0.5 percentage points in Ukraine. The returns to schooling estimates based on earnings actually received during the last six months, which is available only for Ukraine, are close to the estimates obtained using contractual wage. In both countries, taking the hourly wage rate as a dependent variable produces higher estimates of returns to schooling than the baseline estimates. Several additional variables, which are not available for both countries and for all years, are included in the earnings functions to check the sensitivity of the results to the inclusion of these variables. For Russia, we find that including industry dummies has practically no effect on the rates of returns to schooling but adding district fixed effects reduces the estimates significantly by up to 2.1 percentage points.[17]

Family background variables are often used to control for unobserved ability or as an instrument to correct for the possible endogeneity of schooling due to measurement error and omitted ability variables, as Card (1995) and Ashenfelter and Zimmerman (1997) discuss. Because the downward bias resulting from measurement error is often bigger than the upward bias due to omitted ability, OLS estimates are typically lower than instrumental variables (IV) estimates as Angrist and Krueger (1991) and Card (2001) demonstrate.[18] Unfortunately, only the Ukrainian survey has information on parental education and occupation. Nonetheless, the family background effect on Russian returns to schooling is unlikely to be different. We first take parental education and occupation as control variables and obtain the fairly standard result that the estimates of schooling returns are smaller. Next we use family background together with age and age squared as instruments for years of education and labor market experience.[19] Table 4 reports the IV estimates of returns to schooling; returns increase from 7.3% in 1986 to 9.2% in 1991 and then to 12.0% in 2002. These numbers are considerably larger than the corresponding OLS estimates, which is consistent with studies for other transition economies by Heckman and Li (2003) for China and Filer et al. (1999) for the Czech and Slovak Republics. To summarize, although returns to schooling are somewhat sensitive to the choice of variables and specifications, the overall finding of the divergence in rates of returns between Russia and Ukraine should not be affected.

Conceptually, our estimates measure gross monetary returns to schooling. Perhaps low Ukrainian monetary returns are compensated by higher non-monetary benefits to schooling or by lower direct educational costs, which would imply a higher net value of education. For example, more schooling may lead to a lower probability of becoming unemployed.

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[17] The district effect on returns to schooling is unlikely to be different in Ukraine even if this information were available.

[18] Card (2001) discusses several other explanations for larger IV estimates, including unobserved differences between the treatment and comparison groups, specification searching, and heterogeneity in returns to schooling.

[19] The Sargan test cannot reject the validity of these instruments at any reasonable significance level.

Simple descriptive statistics confirm this statement for both countries but the relative difference in unemployment rates between university and secondary school graduates is much higher in Russia than in Ukraine. In 2002, the unemployment rate for individuals with a university degree was 4.5% in Russia and 8% in Ukraine while, for individuals with a secondary education, the unemployment rate was 9.1% in Russia and 12.3% in Ukraine (ILO, 2004; Goskomstat, 2003). Therefore, in Russia educated worker benefit from schooling by having a lower probability of being unemployed than their Ukrainian counterparts.

The direct costs of education are often omitted in conventionally measured returns to schooling. Hypothetically, gross returns may be different but net returns similar if direct costs are significantly lower in Ukraine than in Russia. We follow Fleisher et al. (1996) in computing net returns to schooling as $\beta_{SCH}/(1+\alpha)$ where α is the ratio of the direct costs of education, e.g. tuition, to the indirect costs, i.e. forgone earnings. Specifically, we calculate α as follows:

$$\alpha = \frac{\sum_k T_k S_k (N_k/N)}{W(1-U)}, \tag{4}$$

where N_k/N is the share of students enrolled in school type k, e.g., universities, professional secondary schools, and general secondary schools, T_k is annual tuition fees, S_k is the share of students who pay for their education, W is annual earnings, and U is unemployment rate. In 2002, the computed value of α is 0.063 in Russia and 0.081 in Ukraine.[20] Hence, to obtain net returns to schooling, the estimated coefficient on years of schooling must be multiplied by 0.941 in Russia and by 0.925 in Ukraine. This correction hardly changes the magnitude of the gap in returns to schooling between the two countries and does not influence any of our conclusions.

4. Sources of differences in returns to schooling

Although we cannot establish a causal link between the speed of reforms and returns to schooling, we can investigate the driving forces behind the considerable divergence in returns to schooling, in particular returns to higher education, in Russia and Ukraine.[21] Figure 3 shows kernel density estimates for log wages of university- and secondary school-educated workers for Russia and Ukraine in 1985/1986 and 2002. The densities for university graduates are clearly to the right of the densities for secondary-school graduates in both countries. The estimated mean university gap of wages is much higher in 2002 than in 1985/1986 in both countries; in addition, the gap is higher in Russia at 0.439 than

[20] These calculations are based on the following data in 2002: $W^{\text{RUS}} = 53844$ rubles, $W^{\text{UKR}} = 4512$ hryvnyas, $U^{\text{RUS}} = 8.6\%$, $U^{\text{UKR}} = 10.1\%$, $T^{\text{RUS}} = (22{,}662, 11{,}475, 5675)$ rubles, $T^{\text{UKR}} = (1985, 1100, 900)$ hryvnyas, $S^{\text{RUS}} = (50.9\%, 37\%, 4\%)$, $S^{\text{UKR}} = (59\%, 5.5\%, 5.6\%)$, $N^{\text{RUS}} = (5947.5, 2585.5, 18440)$ thousands of students, $N^{\text{UKR}} = (2269.8, 502.5, 6350.1)$ thousands of students in universities, professional secondary schools, and general secondary schools, respectively (Goskomstat, 2002a, 2002b; Derzhkomstat, 2002; Verhovna Rada of Ukraine, 2004; Ukrainian Ministry of Education, 2004).

[21] We are unaware of any method that decomposes the cross-country differences in returns to schooling as conventionally measured. In this section, we focus on the differences between workers with university and secondary-school education instead of using continuous years of schooling.

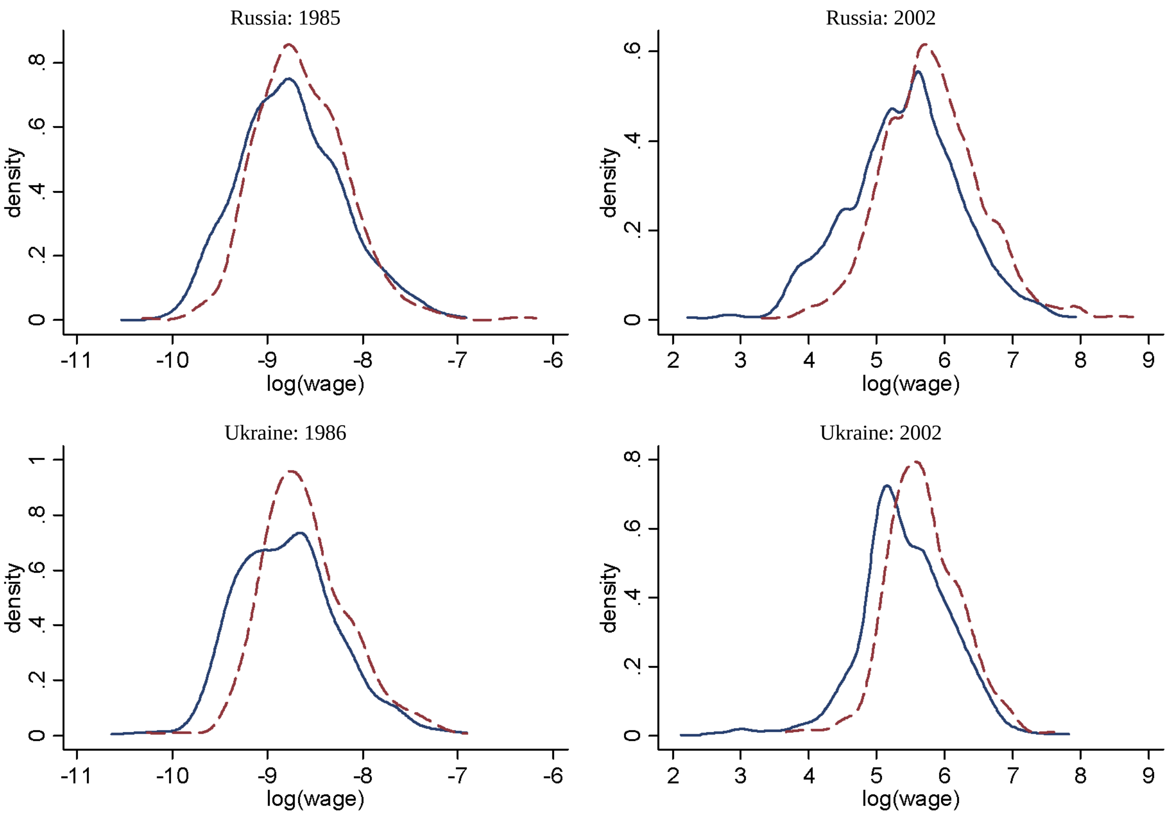

Fig. 3. Kernel density estimates of log wages, Russia and Ukraine. *Notes:* In all graphs, broken and solid lines correspond to workers with university and secondary-school education, respectively. Wages are rescaled so that the mean log wage in Russia is equal to the mean log wage in Ukraine for each year. The plotted densities use the Parzen kernel with bandwidth of 0.45.

in Ukraine at 0.271 in 2002. The difference between the two distributions can be interpreted as a measure of the university wage premium over a secondary-school diploma and specified as:

$$\begin{aligned}
\Delta_k \ln w_t^{\mathrm{RUS}} &\equiv Q_k\big(\ln w_{h,t}^{\mathrm{RUS}}\big) - Q_k\big(\ln w_{s,t}^{\mathrm{RUS}}\big) \\
&= Q_k\big(X_{h,t}^{\mathrm{RUS}}\beta_h^{\mathrm{RUS}} + \varepsilon_h^{\mathrm{RUS}}\big) - Q_k\big(X_{s,t}^{\mathrm{RUS}}\beta_s^{\mathrm{RUS}} + \varepsilon_s^{\mathrm{RUS}}\big), \\
\Delta_k \ln w_t^{\mathrm{UKR}} &\equiv Q_k\big(\ln w_{h,t}^{\mathrm{UKR}}\big) - Q_k\big(\ln w_{s,t}^{\mathrm{UKR}}\big) \\
&= Q_k\big(X_{h,t}^{\mathrm{UKR}}\beta_h^{\mathrm{UKR}} + \varepsilon_h^{\mathrm{UKR}}\big) - Q_k\big(X_{s,t}^{\mathrm{UKR}}\beta_s^{\mathrm{UKR}} + \varepsilon_s^{\mathrm{UKR}}\big),
\end{aligned} \tag{5}$$

where $Q_k(x)$ denotes the kth percentile of variable x, while h and s stand for higher and secondary education, respectively. Figure 4 plots the difference in log wages between the two groups of workers at various percentiles of the distributions. Consistent with the estimates of returns to schooling from quantile regressions, the returns to a university degree are greater in Russia than in Ukraine in 2002. In addition, the university wage premium is generally decreasing with percentiles except for the upper tail in Russia in 2002 and it is largest at the lower percentiles in both countries.

To examine how imported Russian characteristics, i.e., Xs, βs, and εs, could have changed the returns to a university degree in Ukraine, we construct counterfactual distribu-

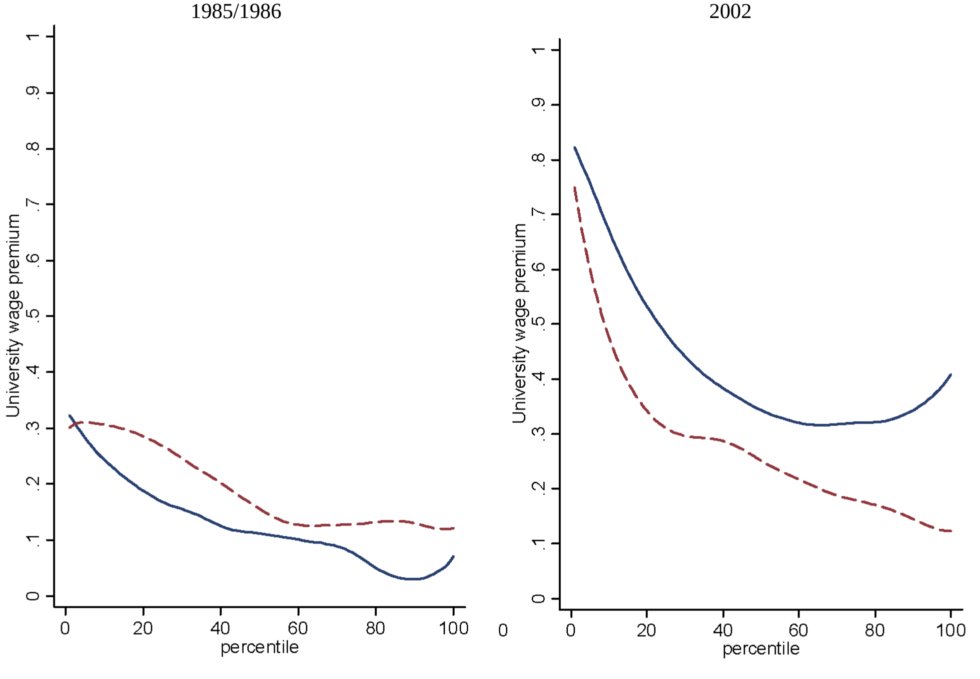

Fig. 4. Distribution of university wage premium, Russia and Ukraine. *Note*. In all graphs, broken and solid lines correspond to Ukraine and Russia, respectively. The university wage premium at a given percentile is defined as the difference in mean wages between university and secondary school-educated workers at a given percentile. A locally weighted regression (lowess) with bandwidth 0.5 is used to smooth percentile estimates in the figure.

tions of log wages for workers with university- and secondary-school education in Ukraine using the distributions of Russian characteristics, returns to these characteristics, and unobservables. These counterfactual distributions provide an estimate of the distributions of Ukrainian log wages that would have prevailed if Ukraine had the same features as Russia. Using actual and constructed wage distributions, we compute and compare actual and counterfactual university wage premia to find the contributions of observable and unobservable characteristics and the returns to cross-country differences in the university wage premium. In general form, the counterfactual university wage premium at each percentile in Ukraine can be written as

$$\Delta_k \ln w_t^{\mathrm{UKR}(m)} \equiv Q_k\left(\ln w_{h,t}^{\mathrm{UKR}(m)}\right) - Q_k\left(\ln w_{s,t}^{\mathrm{UKR}(m)}\right), \tag{6}$$

with m indicating the number of the counterfactual that will be described below.

To construct distributions for the university premia, we first estimate earnings functions from Eq. (3) for each country and for each level of schooling, i.e., university degree and completed secondary school. Then, we generate counterfactual wage distributions following the semiparametric method developed by Juhn et al. (1993), henceforth referred to as JMP. For clarity, we succinctly rewrite Eq. (3) as:

$$\ln w_{r,t}^c = X_{r,t}^c \beta_r^c + \varepsilon_{r,t}^c, \tag{7}$$

where t and c index the time period and country, respectively, $r = \{h, s\}$ denotes the highest attained level of schooling, i.e., higher education with university degree or completed secondary school, w is monthly contractual wages after taxes at the primary job, $\boldsymbol{X}$ is a set of observable characteristics of individuals and firms, and ε is a stochastic error term that absorbs unobservable characteristics of individuals. The coefficients β_r^c can be interpreted as prices for various observable characteristics of workers. In Appendix Table A.3, we present the estimates of Eq. (7) for university and secondary school graduates by country in 2002.

For each level of schooling, we construct four counterfactual wage distributions. First we take Russian observables, Ukrainian prices, and Ukrainian unobservables to yield $\ln w_{r,t}^{\mathrm{UKR}(1)} = X_{r,t}^{\mathrm{RUS}} \beta_r^{\mathrm{UKR}} + \varepsilon_r^{\mathrm{UKR(RUS)}}$. Second, we use Ukrainian observables, Russian prices, and Ukrainian unobservables to generate $\ln w_{r,t}^{\mathrm{UKR}(2)} = X_{r,t}^{\mathrm{UKR}} \beta_r^{\mathrm{RUS}} + \varepsilon_r^{\mathrm{UKR}}$. Third, we consider Ukrainian observables, Ukrainian prices, and Russian unobservables to specify $\ln w_{r,t}^{\mathrm{UKR}(3)} = X_{r,t}^{\mathrm{UKR}} \beta_r^{\mathrm{UKR}} + \varepsilon_r^{\mathrm{RUS(UKR)}}$. Fourth, we have Ukrainian observables, Russian prices, and Russian unobservables resulting in $\ln w_{r,t}^{\mathrm{UKR}(4)} = X_{r,t}^{\mathrm{UKR}} \beta_r^{\mathrm{RUS}} + \varepsilon_r^{\mathrm{RUS(UKR)}}$. The counterfactual unobservables are computed nonparametrically using the JMP method. Specifically, $\varepsilon_{r,t}^{\mathrm{RUS(UKR)}} = F_{\mathrm{RUS},r}^{-1}(F_{\mathrm{UKR},r}(\varepsilon_{r,t}^{\mathrm{UKR}} \mid X_{r,t}^{\mathrm{UKR}}))$, where $\varepsilon_{r,t}^{\mathrm{RUS(UKR)}}$ represents Russian counterfactual residuals corresponding to an Ukrainian individual having a level of schooling r at period t conditional on characteristics $\boldsymbol{X}$, $\varepsilon_{r,t}^{\mathrm{UKR}}$ is the actual Ukrainian residual, F_{RUS}^{-1} denotes the inverse cumulative distribution of Russian residuals, and F_{UKR} denotes cumulative distribution of Ukrainian residuals. The formula for Ukrainian counterfactual residuals is the reverse, namely

$$\varepsilon_{r,t}^{\mathrm{UKR(RUS)}} = F_{\mathrm{UKR},r}^{-1}\left(F_{\mathrm{RUS},r}\left(\varepsilon_{r,t}^{\mathrm{RUS}} \mid X_{r,t}^{\mathrm{RUS}}\right)\right).$$

The difference between the counterfactual wage distribution for the individuals with a university degree and the counterfactual wage distribution for individuals with completed secondary education is the counterfactual university wage premium. Figure 5 plots actual and counterfactual distributions of university wage premium for 2002. The area between the actual university wage premium in Russia and its counterfactual premium in Ukraine, i.e., the distance between the two distributions, can be used in assessing the relative contribution of each factor to the observed differences in returns to a university education. More important factors should bring the Ukrainian counterfactual distributions closer to the Russian actual distribution. To quantify the relative importance of each factor, we take the following measure of the distance between the actual and counterfactual (m) distributions in period t (d_{tm}):

$$d_{tm} = \frac{1}{100}\sum_{k=1}^{100}\left|\Delta_k \ln w_t^{\mathrm{RUS}} - \Delta_k \ln w_t^{\mathrm{UKR}(m)}\right|. \tag{8}$$

If the distributions of wage premium coincide, $d_{tm} = 0$. The larger is the value of d_{tm}, the larger is the difference between the distributions and the smaller is the contribution of the

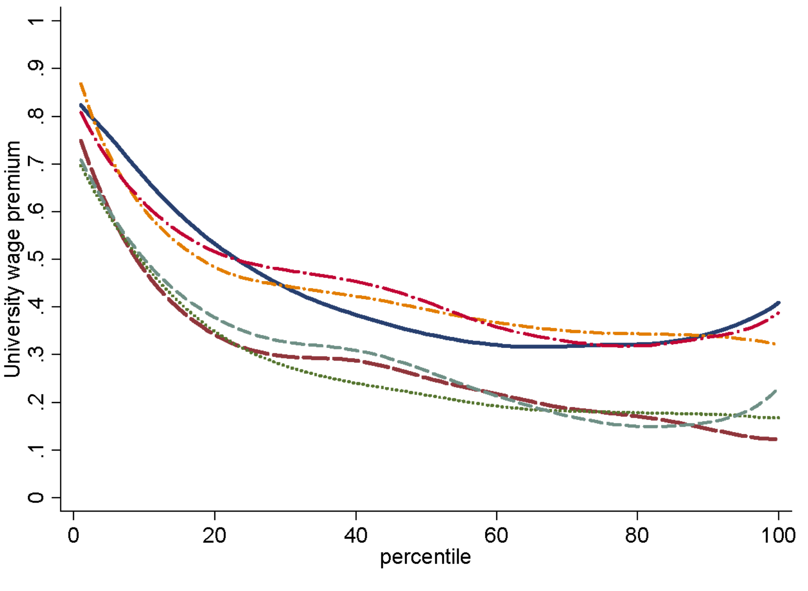

Fig. 5. Actual and counterfactual distributions of university wage premium, 2002. *Note*. The thick solid line is the actual Russian distribution of university wage premium. The thick long dash line is the actual Ukrainian distribution. The bottom dotted line is first counterfactual with Russian observables. The dash/dot line is the second counterfactual with Russian prices. The dash line is third counterfactual with Russian unobservables. The top long dash/dot line is the fourth counterfactual with Russian prices and unobservables. The university wage premium at a given percentile is defined as the differences in mean wages between workers with university and secondary-school education workers at a given percentile. A locally weighted regression (lowess) with bandwidth 0.5 is used to smooth percentile estimates in the figure.

Table 5
Actual and counterfactual distributions of university wage premium, 2002

	Mean	d_{tm}	Selected percentiles				
			10	25	50	75	90
Actual university wage premium							
Russia	0.439	–	0.628	0.382	0.337	0.319	0.379
Ukraine	0.271	0.154	0.470	0.287	0.333	0.182	0.182
Counterfactual university wage premium for Ukraine							
Counterfactual 1	0.277	0.158	0.458	0.290	0.235	0.186	0.180
Counterfactual 2	0.420	0.073	0.476	0.444	0.399	0.356	0.360
Counterfactual 3	0.277	0.138	0.499	0.303	0.303	0.136	0.117
Counterfactual 4	0.427	0.069	0.554	0.485	0.417	0.311	0.311

Notes: (i) Counterfactual 1 corresponds to Russian characteristics and Ukrainian prices for observable and unobservable characteristics. (ii) Counterfactual 2 corresponds to Ukrainian characteristics, Russian coefficients (prices) and Ukrainian unobservable characteristics. (iii) Counterfactual 3 corresponds to Ukrainian characteristics, Ukrainian coefficients (prices) and Russian unobservable characteristics. (iv) Counterfactual 4 corresponds to Ukrainian characteristics, Russian coefficients (prices) and Russian unobservable characteristics (residuals). (v) The distance between the actual Russian distribution and counterfactual Ukrainian distributions is computed from Eq. (8) and denoted d_{tm}.

corresponding factor. Table 5 displays the distance measures along with the key percentiles and means of actual and counterfactual university premia for 2002.[22]

Since Russia and Ukraine have similar observable characteristics, the differences in observable characteristics should contribute very little to explaining the differences in the university wage premium in all years. As Table 5 indicates, Russian observable characteristics are rewarded according to Ukrainian pricing schedules, i.e., counterfactual 1, the mean university premium and the distance measure barely change relative to the actual university premium in Ukraine. In contrast, if Russian slopes are used to price Ukrainian observable characteristics, i.e., counterfactual 2, the mean university premium increases sharply from 27.1 to 42% and the distance between the two distributions shrinks considerably. However, changes in prices do not increase the university premium uniformly. The highest increase in the premium is found in top percentiles, e.g., 18 percentage points at the 90th percentile, and the gain is generally increasing with percentiles.

Because the mean of the counterfactual residuals is close to zero, counterfactual 3 does not change significantly the mean university premium in Ukraine. However, unobservables tend to decrease the premium in the right tail of the distribution and increase it in the left tail thus making its shape closer to the actual university premium in the Russian data. This change reduces the distance between the two distributions, although not significantly. Combining labor force composition of Ukraine with Russian prices for both observable and unobservable characteristics brings the counterfactual wage distributions, i.e., counterfactual 4, even closer to the actual distribution of Russian returns. In summary, the differences in pricing schedules for observed characteristics play a dominant role in explaining differences in university wage premium between Ukraine and Russia.

[22] Results for earlier years are similar to those for 2002. Therefore, we do not report them for reasons of space but they are available upon request.

5. Conclusion

In this paper, we estimate and compare returns to schooling for two countries that belonged to the former Soviet Union and inherited similar institutions and starting conditions, namely Russia and Ukraine. We use the institutional comparability between them and the definitional comparability between two household surveys to examine the cross-country differences in returns to schooling from 1985 to 2002. Our key finding is that, after the breakup of the Soviet Union, returns to schooling diverged significantly between Russia and Ukraine. In 2002, the estimated returns to schooling are two times less in Ukraine at 4.5% than in Russia at 9.2%. We show that this result is remarkably robust to modifications in econometric specifications, definitions of variables, and weighting schemes. Furthermore, we show that the divergence is present not only in the average returns to schooling but also in the distributions of returns to schooling. To investigate the factors responsible for this disparity in returns to schooling, we apply semiparametric methods to construct counterfactual wage distributions for workers having university and secondary-school education. We assess the changes in the university wage premium in response to changes in observable characteristics, prices, and residuals. The labor forces in both countries exhibit similar educational composition and other characteristics during the pre-reform and reform periods. Hence, as expected, we find that the difference in observable characteristics contributes little to observed differences in the university premium across countries.

We conclude that the differences in returns to schooling are unlikely to be supply driven. We also conclude that cross-country differences in unobservable characteristics do not contribute significantly to explaining the differences in returns to schooling. When the Russian unobservable characteristics are combined with Ukrainian observable characteristics and Ukrainian prices, the shape of the distribution of the university premium in Ukraine becomes closer to the one in Russia but the distance between the two distributions remains significant. Perhaps a common history, active migration of families between Russia and Ukraine, similar human capital and abilities, the same preferences for higher education, and shared institutional and organizational practices yield similar unobservable characteristics in the two countries. In contrast, differences in prices of observable characteristics play a critical role. If Ukrainian workers had been rewarded according to Russian pricing schedules, the educational premium would be comparable to that in Russia. Although the reason for these price differences requires further study, we conjecture that the lower demand for educated labor, more limited labor mobility, higher separation costs, and the larger role played by trade unions in Ukraine are the most likely explanations for the differences in returns to schooling between the two countries.

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Appendix Table A.1
Definitions and sources of main variables

Variable	Russia	Ukraine
Wage	• 1998–2002: Average monthly wage after taxes at the primary job, regardless of whether it was paid on time or not, for employees; monetary and in-kind payments actually received after taxes at the primary job in the last 30 days for self-employed. • 1994–1996: Imputed the contractual wage as the ratio of the total wage debt to the number of monthly wages owed; monetary and in-kind payments actually received after taxes at the primary job in the last 30 days for employees without wage arrears and self-employed. • 1985–1990: Average monthly wage.	Monthly contractual wage after taxes at the primary job in December of the corresponding year. All wages are converted into hryvnyas.
Adjusted years of schooling	Education status is converted into a continuous variable representing adjusted years of schooling. For consistency with ULMS, adjusted years of schooling were taken as 4 for 1–6 grades, 8 for 7–9 grades, 10 for 10–12 secondary school grades, 9 for a vocational non-secondary school diploma, 11.5 for a vocational secondary school diploma, 13 for a technical school diploma and incomplete higher education, 15 for a diploma of specialist, and 18 for a PhD degree.	Education status is converted into a continuous variable representing adjusted years of schooling. Adjusted years of schooling were taken as 4 for 1–6 grades, 8 for 7–9 grades, 10 for 10–12 secondary school grades, 9 for a vocational non-secondary school diploma, 11.5 for a vocational secondary school diploma, 13 for a technical school diploma and incomplete higher education, 14 for a bachelor degree, 15 for a diploma of specialist, 16 for a master degree, and 18 for a PhD degree. Educational histories are used to compute adjusted years of schooling for previous years. The same definitions are used to compute adjusted years of schooling of parents.
Actual years of schooling	Total number of years in a school including part-time schools, evening schools, and courses by correspondence; available for 1995–2002.	Total number of years in a school including part-time schools, evening schools, and courses by correspondence; available for 2002.
Potential labor market experience	Age minus years of schooling minus 6.	Age minus years of schooling minus 6.
Tenure	Number of years since an individual started the primary job.	Number of years since an individual started the primary job.
Weekly hours of work	2002: Average hours in the usual work week at the primary job.	2002: Hours per week an individual usually works at the primary job; not available for other years.
Hours of work missing	= 1 if hours of work is missing.	= 1 if hours of work is missing.
Parents' occupations	N/A	Dummy variables for a manual non-agricultural worker, a collective farmer/agricultural worker, a non-manual worker.

(*continued on next page*)

Appendix Table A.1 (*continued*)

Variable	Russia	Ukraine
Foreign	= 1 if primary employer is owned or co-owned by foreign firms or foreign individuals.	= 1 if primary employer is domestically owned with some foreign capital or foreign-owned (including international organizations).
Private	= 1 if primary employer is owned or co-owned by Russian private firms or Russian individuals (with no foreign participation); or if an individual is self-employed.	= 1 if primary employer is a privatized enterprise, a newly established private enterprise, or a cooperative (with no foreign participation); or if an individual is self-employed.
State	= 1 if primary employer is owned by state.	= 1 if primary employer is a budgetary organization, a state enterprise, a local municipal enterprise, a state farm, or a collective farm.

Appendix Table A.2
Employment distribution by the level of schooling

Panel A: Russia

Schooling level	1985	1990	1996	1998	2000	2002
Secondary school (0–6 grades)	8.04	4.24	0.94	0.57	0.31	0.21
Secondary school (7–9 grades)	15.51	12.45	8.97	7.76	6.96	6.98
Vocational school with no high school diploma	5.66	5.35	4.06	4.31	3.94	3.43
Secondary school (10–12 grades)	23.32	24.19	24.51	23.67	23.68	23.31
Vocational school with high school diploma	10.68	13.67	15.08	16.48	18.23	17.52
Technical school	20.88	22.87	24.79	25.58	25.98	25.69
University	15.33	16.58	20.79	20.67	20.18	22.02
Graduate school	0.58	0.64	0.86	0.96	0.72	0.84
N	4111	3776	3497	3332	3169	3341

Panel B: Ukraine

Schooling level	1986	1991	1997	1998	2000	2002
Secondary school (0–6 grades)	6.56	2.95	1.02	0.67	0.61	0.46
Secondary school (7–9 grades)	12.77	10.39	6.75	5.89	4.20	4.16
Vocational school with no high school diploma	7.19	7.64	7.69	7.66	7.71	7.63
Secondary school (10–12 grades)	27.03	27.28	25.46	25.26	23.99	24.40
Vocational school with high school diploma	11.62	14.56	15.12	16.03	17.41	17.44
Technical school	21.56	22.54	26.17	26.36	26.76	25.71
University	12.83	14.23	17.69	18.02	19.11	20.02
Graduate school	0.46	0.40	0.10	0.11	0.20	0.18
N	4191	3528	2946	2812	2925	3289

Notes: (i) The sample is restricted to respondents aged 15 to 59 years with non-missing values for the variables used in the basic Mincerian wage function. (ii) The sample weights are applied for 1985/1986 and 1990/1991 in both countries.

Appendix Table A.3
Earnings functions for university and secondary school graduates, OLS, 2002

	Ukraine		Russia	
	University (1)	Secondary school (2)	University (3)	Secondary school (4)
Female	0.244***	0.350***	0.391***	0.446***
	(0.045)	(0.033)	(0.051)	(0.042)
Experience (years)	0.018*	0.014**	0.041***	0.024***
	(0.009)	(0.006)	(0.010)	(0.009)
$\text{Experience}^2/1000$	−0.424*	−0.357**	−0.979***	−0.612***
	(0.239)	(0.146)	(0.268)	(0.203)
Capital	−0.037	0.355***	0.482***	0.655***
	(0.074)	(0.065)	(0.084)	(0.090)
Tenure (years)	0.005	0.008	0.019*	−0.004
	(0.008)	(0.007)	(0.010)	(0.008)
$\text{Tenure}^2/1000$	−0.104	−0.222	−0.538	0.162
	(0.258)	(0.219)	(0.352)	(0.283)
Ownership				
Private	0.292***	0.253***	0.417***	0.334***
	(0.062)	(0.042)	(0.057)	(0.047)
Foreign	0.076	0.523***	0.745***	0.677***
	(0.157)	(0.115)	(0.127)	(0.107)
Employer size (no. of persons)				
10–50	−0.049	0.024	−0.067	−0.024
	(0.091)	(0.061)	(0.090)	(0.068)
50–100	0.125	0.012	−0.024	0.205**
	(0.094)	(0.076)	(0.103)	(0.090)
100–500	0.143	0.217***	0.130	0.091
	(0.099)	(0.061)	(0.094)	(0.072)
500–1000	0.215*	0.308***	0.064	0.300***
	(0.124)	(0.072)	(0.118)	(0.094)
>1000	0.370***	0.475***	0.005	0.413***
	(0.099)	(0.064)	(0.109)	(0.087)
Constant	5.224***	4.852***	7.373***	7.017***
	(0.120)	(0.087)	(0.120)	(0.095)
N	663	1375	753	1345
R^2	0.16	0.19	0.25	0.20

Notes: (i) Dependent variable is the log of monthly contractual wages after taxes at the primary job. (ii) Robust standard errors are in parentheses. (iii) Sample weights are applied in Russia and the sample is restricted to persons with age of 15 to 59 years. (iv) The omitted categories are 1–10 for employer size and state for ownership. (v) Two dummy variables for missing employer size and missing ownership are included but their coefficients are not reported.

* Significance at the 10% level.
** Idem., 5%.
*** Idem., 1%.

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