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Ownership and Wages: New Evidence from Linked Employer-Employee Data in Hungary, 1986-2003

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Abstract

How do state, domestic private, and foreign ownership of firms differ in average wages and skill differentials? We address these questions using linked employer-employee panel data containing 1.4 mln worker-year observations for 21,381 firms, of which more than 6000 change ownership type, from 1986 to 2003 in Hungary. Our analysis shows a sharp rise in wage inequality that is strongly associated with ownership changes, particularly the shift to foreign ownership. OLS estimates imply a substantial positive wage differential in foreign firms and a negative differential in domestic private, relative to state enterprises, but regressions including firm fixed-effects and trends imply no foreign and only small domestic differentials. Foreign firms exhibit larger educational and occupational differentials than either state or domestic private firms, but the results are again sensitive to the inclusion of firm fixed effects and trends.

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1. Introduction

The level and structure of wages in the transition economies of Eastern Europe have changed dramatically in the 15 years since the collapse of central planning. Real wages tended to decline rapidly in the first few years of transition and to rise strongly more recently, while both overall inequality and estimates of wage differentials associated with schooling show large increases in every country where they have been studied.[1] At the same time, the economies of the region have experienced massive organizational changes, most prominently large-scale privatization and opening to the global economy, including foreign direct investment.

These rapid changes provide a useful context for investigating the relationship between firm ownership and the level and structure of wages. The transfers from the state to new domestic and foreign owners took place not only quickly but broadly across nearly all sectors. The tightly controlled wages of the centrally planned systems were abruptly liberalized, permitting organizations to set their own wages and to increase skill differentials, which had tended to be compressed under socialism (e.g., Kornai, 1992). But how these changes might be related is unclear a priori. If firms maximize profits, labor markets are perfectly competitive, there are differences in fixed costs of employment, and the wage equals the full value of the job to workers, then wages should be correlated with ownership only through compositional differences in types of employees. Shifts in labor demand may lead to temporary wage differentials for the same type of worker, but these should disappear as workers move from lower to higher return activities. However, if ownership is associated with different objectives, fixed

[1] Commander and Coricelli (1995) and World Bank (2005) document average real wage changes in a number of transition economies. Studies of schooling differentials include Vecernik (1995), Rutkowski (1996), Brainerd (1998), Chase (1998), Flanagan (1998), Filer et al. (1999), Kertesi and Kollo (2002), Campos and Jolliffe (2003), Sabirianova Peter (2003), Andren et al. (2005), Fleisher et al. (2005), Gorodnichenko and Sabirianova Peter (2005), and Münich et al. (2005a, b).

costs, fringe benefits, or other work conditions, then differences in wages across these types may persist even beyond the time required for workers to overcome mobility frictions.

In this paper, we estimate the relationship between wages and ownership using linked employer-employee panel data for Hungary. Hungary is a particularly apt country for the analysis, not only because it underwent sweeping ownership changes, similar in this sense to some of its neighbors, but also because its privatization policies tended to result in ownership structures more akin to those in market economies, with more outside investor control and with much more foreign involvement. Moreover, the available data are exceptional in size and quality. The data include observations on some 1.4 mln worker-years at 21,381 employers that we follow over a long time period, from 1986 to 2003. The worker characteristics in the data are useful for controlling for the composition of employment at each firm and for estimating skill differentials, and the firm-side information permits us to measure ownership changes precisely and to control for some types of selection bias into ownership type. However, the data allow us to distinguish only three types of ownership: state, domestic private, and foreign. They also do not enable us to follow individual workers over time, nor do they include information on working hours, nonmonetary benefits, and other work conditions. We thus cannot control for unobserved differences across workers, nor can we rule out the possibility that observed wages reflect compensating variations with respect to differences along other dimensions of the employer-employee relationship.

Nevertheless, these data help overcome a number of drawbacks in previous research. Studies relying on firm-level data usually have small samples, very short time series, and no worker characteristics, and they sometimes lack a comparison group. Identification may depend on observing ownership changes, but few studies analyze the effects of privatization on wages.[2] Haskel and Szymanski (1993) is the earliest systematic study, and it analyzed 14 British publicly owned companies, of which only four were actually privatized. Martin and Parker (1997) study 14 large British privatizations, while Kikeri (1998) and Birdsall and Nellis (2003) summarize a number of case studies and small sample surveys of privatization effects on labor in several developing economies. La Porta and Lopez-de-Silanes (1999) analyze 170 privatized firms in Mexico, although the post-privatization information is limited to a single year. The small sample size problem is overcome in Brown et al. (2005), who study nearly comprehensive panels of manufacturing firms in Hungary, Romania, Russia, and Ukraine, finding a zero or small negative effect of privatization.[3] But a fundamental problem with all of this work using firm-level data is the inability to measure worker characteristics and thus to control for composition of the workforce, particularly if changes in composition are correlated with changes in ownership.

A similar problem is evident with most studies of relative wages at foreign-owned firms. For example, Feliciano and Lipsey (1999) study wage differentials between foreign and domestically owned establishments in the United States. Aitken et al. (1996) analyze the same topic, but extend the analysis with wage spillovers between foreign and domestic firms. Conyon et al. (2002) study wage changes following foreign acquisitions in manufacturing firms in the United Kingdom. Finally, Lipsey and Sjöholm (2004) study these wage differentials in Indonesian manufacturing, although in this case they do control for the composition of workforce at the firm level. All these studies find a wage premium in foreign firms.

[2] The lack of research on the wage impact of privatization contrasts with the large literature on firm performance, already the subject of multiple survey articles (e.g., Megginson and Netter, 2001; Djankov and Murrell, 2002).

[3] A related line of research analyzes effects of all types of ownership change on wages: e.g., Lichtenberg and Siegel (1990) on leveraged buyouts, Gokhale et al. (1995) on hostile takeovers, and McGuckin and Nguyen (2001) on mergers and acquisitions. Our data do not contain information on all ownership changes, but only on transitions between state, domestic private, and foreign ownership types, which are thus our focus in this paper.

However, a second problem, equally serious, is that most studies do not account for ownership selection effects. If firms experiencing an ownership change are not randomly selected with respect to their wage behavior and the researcher does not take this into account, the estimated effect of ownership change will generally be biased. Indeed, some recent studies demonstrate this possibility.[4]

Instead of using firm-level data, another category of research has employed individual data that include information on employer ownership as well as wages. The sizable literature on public-private wage differentials is surveyed by Gregory and Borland (1999), but in the Western context this amounts to an analysis of interindustry differentials with little possibility of taking into account unobserved differences in ownership types that are correlated with wages. Concerning foreign wage differentials, Peoples and Hekmat (1998) carry out an analysis for the US, but they use only industry-level ownership information. In the transition context, Brainerd (2002) estimates wage effects of Russian mass privatization using worker-level data but in a cross-section approach not distinguishing between privatized and new private firms. On the wage structure, Münich et al. (2005a) report no difference in the wage impact of education for Czech women working in state and privately owned enterprises. But this result is unusual: Rutkowski (1996) and Keane and Prasad (2004) find that private ownership increases the education coefficient in Poland; Kertesi and Kollo (2002) find a higher education coefficient under foreign compared to domestic ownership in Hungary; and Andren et al. (2005) find a positive private premium and larger schooling differential in private firms in Romania. A problem with these studies is possibly inaccurate measures of ownership, which are reported by

[4] Conyon et al. (2002) employ firm fixed effects to study foreign acquisitions in Britain. Almeida (2003) discusses selection of foreign acquisitions, and Brown et al. (2005, 2006) discuss selection in privatization programs.

workers who may not be fully informed about the progress of the privatization process. More importantly, worker-level data do not permit controls for firm selection into ownership type.[5]

The advantages of both firm- and worker-level data can be exploited only if one combines the two data types into linked employer-employee data. But only two previous studies, both of them recent working papers, use linked data for a similar purpose, and both focus on the effects of foreign acquisitions on wages in Portugal: Almeida (2003) finds higher wages in foreign firms, but Martins (2004) reports a negative effect. These studies share the problem, common to most Western data sets, of relatively few ownership changes, so that the ownership effect is identified only on a small sample of firms. In our Hungarian data, by contrast, we observe thousands of ownership changes, including about 5000 involving domestic private ownership and 1100 involving foreign ownership.

We use these data to estimate differences in wage levels and structure by ownership type, recognizing that these differentials are unlikely to equal the causal effects of ownership. First, it is likely that selection of firms and workers into ownership types is nonrandom with respect to unobserved factors, such as quality of the firm or the worker. We exploit the longitudinal structure of the firm side of the data to control for fixed and trending differences across firms, but because we do not know the form taken by the heterogeneity we cannot be sure that these methods fully account for selection bias. Moreover, we cannot control for unobserved heterogeneity at the worker-level. A second issue in interpreting our estimates on domestic private and foreign ownership is that we do not observe wage outcomes in state firms under a counterfactual of no privatization and no liberalization of foreign entry into the Hungarian economy. Indeed, wage behavior of each ownership type may well be influenced by each of the

[5] An identification approach in analyzing wage differentials across sectors examines wage changes of workers who switch sectors (Krueger and Summers, 1988). Our firm fixed effects method below relies on firms switching sectors.

others through labor market interactions. Analyzing such spillover effects could be interesting, but we leave it for future research.

The next section describes the construction of the employer and employee components of our data, and how we link them into a single database. In Section 3, we briefly explain the changes in the ownership structure during the period studied and summary statistics for all variables. We also provide some initial analysis of the evolution of wage levels, inequality, and some aspects of differentials by worker characteristics. Section 4 describes regression estimates of the impact of ownership on the level and structure of wages, including specifications that control for selection bias into ownership type based on firm-specific time-invariant and time-trending heterogeneity. An important issue in estimating such impacts is the appropriate unit of analysis, and we provide some comparisons of results where the observation is a worker-year with others where the observation is a firm-year. Our data measure wages at both levels, but the worker-year observations permit us to analyze worker heterogeneity in wages and to control for worker characteristics, while the firm-year approach is more closely aligned with our variable of interest, firm ownership. Section 5 concludes with a summary and suggestions for further research.

2. Data Sources and Sample Construction

We study a linked employer-employee dataset from two sources. The first is the Hungarian Wage Survey, which is conducted by the Hungarian Statistical Office to gather information on individual worker characteristics and wages. The Wage Survey was carried out in 1986, 1989, and annually since 1992, with the last available round in 2003. Our analysis thus uses information on workers from 1986, well before the transition started (in 1990), until 2003, the year just prior to European Union accession. Until 1995, the sampling frame each year includes every firm in Hungary with at least 20 employees; after 1995, the size threshold for inclusion is 10 employees and a random sample of smaller firms is also included. To maintain consistency across years, we restrict attention to firms with at least 20 employees in at least one year.

From this sampling frame of firms a sample of workers is selected and included in the Wage Survey database. In 1986 and 1989, workers were selected by using a systematic random design with a fixed interval of selection. From 1992 the sample design changed: production workers were selected if born on the 5th or 15th of any month, while non-production workers were chosen if born on the 5th, 15th, or 25th of any month. This selection procedure thus provides a random sample of workers within firms and includes, on average, about 6.5 percent of production workers and 10 percent of non-production workers. The data also provide weights to adjust for the oversampling of nonproduction workers.[6] Another potential problem is that a firm drops out from the sample if no employee was born on the relevant dates, and the probability of being included therefore increases with the size of the firm.[7] In addition, for small firms which are included in the sample, the whole workforce may be represented by very few workers, which could result in measurement error. Finally, the data do not provide identification codes for workers, so it is not possible to track them across years.

The Wage Survey data includes some firm characteristics, which we use to link to our second data set, from the Hungarian Tax Authority.[8] This database consists of annual firm-level information between 1992 and 2003 on every firm that used double-entry bookkeeping. The

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[6] Using these weights does not affect the results, and we report the unweighted results in this paper.
[7] For example, a firm with 20 production workers will approximately have a probability of 0.11 to be excluded from the sample, while for a similar firm with 100 employees this probability is only 0.012. As one check, we have estimated all equations restricting the sample to employees of firms with more than 100 workers, and the results are qualitatively similar to what we report for the larger sample.

information includes the balance sheet and income statement, the proportion of share capital held by different types of owners, and some basic variables, such as the location and industrial branch of the firm. We use the share capital variables to construct the ownership structure. For the two early years – 1986 and 1989 – the Tax Authority data are not available, and for these years we use the firm information from the Wage Survey; ownership in these years is always state, so the share capital variables are not necessary.

We cleaned firm ownership data extensively, checking for miscoding and dubious changes (e.g., firms that switch back and forth between ownership types). Our procedures also paid a great deal of attention to longitudinal links, for which used a dataset from the Central Statistical Office of Hungary providing information on reregistration and boundary changes. As this dataset is not comprehensive, we also tried to find spurious entries and exits by looking for matches of exits among the entries on the basis of headquarter settlement, county, industry, and employment.

Finally, we focus attention on types of firms in which ownership could conceivably vary between state, domestic private, and foreign. Our selection rule is to include firms only from those industries which had at least 200 firm-years under each of domestic private and foreign ownership in our data. The following industries satisfied this criterion: mining, manufacturing, utilities, construction, wholesale and retail trade, hotels and restaurants, finance, and real estate. Agriculture, education, health, public administration, and the military are excluded.

Table 1 shows the number of workers with full information on characteristics, the number of firms with information on ownership, and the total number of employees in these firms. The

[8] Neither data set contains firm names, exact addresses, or identification codes, and we constructed the links using an exact one-to-one matching procedure for the following variables: county (*megye*), detailed industry, and financial indicators such as sales and profits.

data set we work with is a panel of 21,831 firms linked with a within-firm random sample of almost 1.4 million workers.

3. Evolution of Ownership, Variable Definitions, and Summary Statistics

Compared with its neighbors in Eastern Europe, Hungary began corporate control changes relatively early. Starting with a more relaxed planning regime in 1968, the socialist government gradually permitted state-owned enterprises to operate with increased autonomy, and the decentralization process accelerated during the 1980s (e.g., Szakadat, 1993). Movement of assets out of state ownership began at the very end of the 1980s in the form of so-called "spontaneous privatization," which usually involved spin-offs initiated by managers, who were also usually the beneficiaries, sometimes in combination with foreign or other investors (e.g., Voszka, 1993). After the first free elections in May 1990, procedures became more regularized, involved sales of entire going concerns, and generally relying upon competitive tenders open to foreign participation. Unlike the programs in many other countries, the Hungarian policies did not grant workers significantly preferential prices at which they could acquire shares in their companies, with the exception of about 350 management-employee buyouts. Nor did Hungary carry out a mass distribution of shares aided by vouchers, as was common in most other countries of the region. On the other hand, Hungary was much more open to foreign investors than elsewhere. As a consequence, Hungarian privatization resulted in very little worker ownership, very little dispersed ownership, and high levels of blockholdings by managers and both domestic and foreign investors.[9]

Our database provides the ownership shares of the state, domestic, and foreign owners at the end of each year (the reporting date). We consider a firm to be under domestic private ownership if it is majority private and the domestic ownership share is higher than that of foreign ownership. If the foreign share is larger than the domestic, the firm is foreign-owned, for the purposes of this paper.[10] The evolution of the ownership structure among the firms in our sample is presented in Figure 1, clearly reflecting the early start and the heavy presence of foreign ownership in Hungarian privatization. Although there was only negligible privatization and new private entry by 1989, most of the firms in our sample are majority private by 1992.[11] The share of domestically privatized firms grew steadily until 1998, when almost three-quarters of firms were controlled by domestic owners. Thereafter, it ceased growing and even shrank slightly (because of attrition from the sample). The proportion of foreign-owned firms grows steadily in our sample, reaching 20 percent by 2002. At the same time, the state kept a majority stake in at least 5 percent of firms in each year, thus providing a comparison group for the effects of privatization.

Table 2 shows the incidence of various types of changes in ownership type. The transition process resulted in many more changes from state to private than could ever be observed in a nontransition economy, and the number of changes involving foreign ownership in Hungary are probably the largest that could be found in Eastern Europe. In our data, about 5000 ownership changes involve domestic private ownership, and about 1100 involve foreign ownership. We will exploit these ownership changes when we control for unobserved heterogeneity in estimating wage differentials, as described below.

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[9] Frydman et al. (1993) and Hanley et al. (2002) contain descriptions of the Hungarian privatization process. Earle et al. (2005) study ownership of firms listed on the Budapest Stock Exchange.

[10] This definition has the advantage over definitions that would involve majority ownership that all privatized firms can be categorized as domestic- or foreign-owned. We also estimated specifications using the percentage of shares held by private domestic and foreign owners, with results very similar to those reported below.

[11] These percentages are based on unweighted numbers of firms; if weighted by size, the state-owned sector would be much larger.

The wage variable in our data is gross monthly cash earnings in March, which we have deflated by the annual CPI.[12] Figure 1 shows the evolution of real wages from 1986 to 2003: an initial decline of around 30 percent and subsequent rise of about 45 percent.[13] The steady, substantial growth in the Hungarian real wage since the mid-1990s is unusual among the transition economies, and an interesting question is whether Hungary's relatively rapid privatization and large foreign component may have contributed to this performance. The reliability of the real wage measure is of course strongly influenced by the quality of the deflator (in this case, the CPI), and the large changes in quality and availability of goods suggest caution should be exercised when interpreting these figures. When we estimate wage differences by ownership, however, we include year effects, so our comparisons are not influenced by these measurement problems.

To document changes in wage inequality, Figure 2 plots the variance of ln(wage) for each cross-section of the data. The variance is less than 0.2 in 1986 but it increases substantially until the year 2000, when it reaches 0.46 and then declines to about 0.38. Compared with inequality in the U.S., Hungarian inequality by this measure was lower in the early years, but at its peak exceeds that in the U.S.[14] Thus, the relative rise in Hungarian inequality is much greater.

The degree to which inequality may be associated with ownership types is addressed by a decomposition of the variance of the ln(wage) σ^2 into components attributable to inequality within and across ownership types:

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[12] Most studies of wages in Eastern Europe (and many in Western Europe) analyze monthly rather than hourly or weekly earnings, because of institutional differences such as the custom of reporting wages on a monthly basis, the lower incidence of part-time employment and greater standardization of full-time hours, and the frequent unavailability of hours information (even for production workers). In our data, hours of work are available only for the most recent years, so we cannot analyze changes using them.

[13] To maintain comparability over time, the evolution of the average real wage is estimated as the year effects in a ln(real wage) equation that controls for firm fixed effects.

[14] Figures in Katz and Autor (1999), for example, imply a U.S. variance of ln(weekly wage) for full-time, full-year workers of 0.36 in 1995, rising from 0.25 in 1963 (calculated from Table 1, p. 1475).

$$\sigma^2 = \Sigma_i S_i \sigma_i^2 + \Sigma_i S_i (w_i - w)^2,$$

where S_i is the employment share of ownership type i (state, domestic private, foreign), σ_i^2 is the variance of ln(wage) for type i, w_i is the mean ln(wage) of type i, and w is the mean ln(wage) in the entire sample. Table 3 contains the results from these calculations for 1992-2003. The increase in the variance is mostly accounted for by changes in the first term, the components representing the share-weighted within-ownership variances. The contribution of state sector inequality to the total variance declines sharply, while the domestic private and foreign contributions rise quickly to more than compensate for this decline. Most of these effects are due not to changing variances within ownership types, but instead to higher variance among private firms, especially among those foreign-owned, together with changing shares: the declining state employment share and rising domestic private and foreign shares. The between component, however, contributes relatively little to overall inequality, accounting for only about a third of the overall rise in the variance over this period. In accounting for rising inequality, ownership differences in mean wages appear to be less important than differences in variances.

To begin to analyze systematic factors lying behind the differing variances within ownership types, we turn next to individual worker characteristics. Figure 3 documents the evolution of wage differentials by educational attainment. Four categories of education are used: elementary, vocational, high school, and university.[15] As the figure shows, the vocational premium over elementary education changes little over this period, but the university premium over each of these rises by about 50 percent, and the high school premium by close to 20 percent.

[15] These categories are constructed from more disaggregated variables. "Elementary" refers to 8 or fewer years of schooling, mostly 8 in our data. "Vocational" refers to a short secondary school focused on a trade and involves 10-11 years, usually 11. "High school" refers to completed secondary school, 12 years. "University" refers to any post-secondary schooling, including both *foiskola* (technical higher education involving 3-4 years after high school) and undergraduate and graduate education at universities and institutes (at least 5 years).

At its peak, the university-high school differential exceeds 0.7. Both the level and the changes are larger than those in the U.S. that have been studied so extensively.[16]

Figure 4 shows occupational wage differentials relative to a base category of unskilled manual worker. While the evolution of the occupational wage structure shows relatively little change for some groups (skilled manual and service workers), in general there is substantial widening, with the skilled nonmanual and associate professional log differentials moving in parallel and increasing by about 0.4. While the highest wages are paid to managers, the largest increase in the log differential is for professionals, with a rise from 0.5 in 1986 to 1.2 in 2000.

Table 4 provides calculations of differences in mean wages and characteristics of workers by type of owner, presenting information for 1992 and 2003 – the first and the last year in our panel when each ownership type is present. In both years, the unconditional mean wage is smallest in domestic private firms, largest in foreign-owned firms, and intermediate under state-ownership. Average worker characteristics also vary, however, with higher rates of female and university employment in foreign-owned firms, higher rates of vocational employment in domestic private firms, and higher rates of high school employment under state ownership.[17] Potential experience tends to be lower in foreign-owned firms, a difference that becomes much more pronounced by 2003. The composition of the workforce by occupation also varies considerably, with a much higher rate of employment of professionals under foreign ownership, and a high rate of skilled manual employment in domestic private firms. Such factors likely influence average wage differentials by ownership type and can be taken into account by multivariate analysis.

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[16] For example, Katz and Autor (1999), report college/high school ln(weekly wage) differentials of 0.36 in 1963 and 0.51 in 1995.

In the next step towards such an analysis, Table 5 contains calculations of mean wages by ownership type and educational attainment in 1992 and 2003. For both years and all four educational categories, the ownership types are clearly ranked in wage levels, with foreign highest, state second, and domestic private lowest. The implied wage premium associated with university education, measured relative to either elementary or high school, follows the same order. At this level of analysis, there are clearly large differences among the three ownership types in both the level and the structure of wages they pay. Moreover, it is striking that the wage behavior associated with the two types of private ownership – domestic and foreign – should be much more different from each other than from state ownership.

4. Regression Estimates

To estimate the systematic impact of ownership on wages, we turn to regressions. We are interested not only in controlling for worker characteristics in various combinations – and in assessing the robustness of our results to such controls – but also in attempting to remove some types of selection bias in the determination of ownership type. For example, if state-owned enterprises that already pay higher wages are more likely to be purchased by foreigners (perhaps because of higher unobserved skill, better technology, or indeed for any reason), then the foreign wage premium we have documented may be due to the systematic selection of high wage firms into foreign ownership. And if foreign investors are more likely to acquire firms with greater pay disparities, for instance because of higher premia earned by the university-educated and by managers and professional employees, then again the observed differentials reflect a selection rather than causal effect. The privatization process involving domestic owners may also have

[17] Wages and educational composition for the categories never privatized and eventually domestic and foreign privatized firms are much more similar in 1986 than in Table 2, indicating that the different composition and wages in 1992 are probably due at least partly to privatization.

biases, as politicians, frequently together with employees, choose whether a firm is privatized. Politicians may prefer to retain firms with the worst prospects in state ownership in order to protect workers from layoffs and wage cuts, and the employees themselves may work to prevent privatization in such cases. If the privatization process is corrupt, then exactly the opposite may be true: politicians may prefer to sell the best firms quickly in order to collect bribes.

Of course, we cannot entirely eliminate all possibility of such biases, but the large number of ownership changes together with the longitudinal dimension of our data permit us to at least check whether the differentials implied by our analysis so far are robust to some simple attempts to account for selection bias. For this purpose, we employ methods developed for the evaluation of training programs in the U.S. The first method is the standard correlated effects model that controls for time-invariant unobserved heterogeneity at the firm level. A second is the random growth model, which includes a firm-specific linear time trend.[18] Such a model may be appropriate if, for example, foreign investors are more likely to acquire firms that for some intrinsic reason (unobservable to the researcher, but not caused by ownership) are raising their wages or increasing the premia paid to more highly educated workers. Higher-order parameterizations of heterogeneity are of course possible, but we do not take them into account, and identification of the effect of ownership in our analysis assumes that any other heterogeneity is uncorrelated with either ownership or wages (or wage differentials, when that is our focus). Both of these estimators rely on ownership changes to identify the coefficients of interest; indeed, the random growth model measures changes in the growth rate before and after an ownership change. A resulting disadvantage is that the results pertain to firms that experience such changes,

[18] Ashenfelter and Card (1985) and Heckman and Hotz (1989) use random trend models to evaluate training, while Jacobson et al. (1993, 2005) apply it to the wage effects of job displacement and community colleges. Brown et al. (2005, 2006) use the model to estimate the impact of privatization on employment, wages, and productivity at the

not to the broader sample.[19] Finally, we use some specification tests to evaluate the performance of the estimators, including a variant of the "pre-program test" proposed by Heckman and Hotz (1989).

All equations control for year of observation and region of the establishment. We report standard errors in all cases permitting general within-firm correlation of residuals using Arellano's (1987) clustering method, so that our test statistics are robust to both serial correlation and heteroskedasticity.[20] Standard errors are also adjusted for loss of degrees of freedom in specifications when the data are demeaned and detrended. We first analyze the estimates of the impact of ownership on the wage level and then turn to the impact on wage structure.

Table 6 displays estimates by pooled OLS, firm fixed effects estimations (FE), and firm fixed effects and trends (FE&FT). The first OLS column includes no controls beyond year and region, and the estimates demonstrate that the raw ownership differences are large (-0.26 for domestic and 0.18 for foreign), and they are precisely estimated. Together they imply a large foreign-domestic differential of about 0.44, again highly statistically significant. The next column adds standard variables – education, experience, and gender – to construct a Mincer earnings function, but with little qualitative change in the results: the magnitude of the domestic coefficient declines and the foreign coefficient rises. The little difference between these

firm level. Our paper is the first to our knowledge that uses firm-level trends in any analysis of worker-level wages, and it is the first that uses firm fixed effects in a study of ownership and worker-level wages.

[19] Another potential disadvantage is that these estimators may raise the noise-to-signal ratio, eliminating relevant between-firm variation while exacerbating the effects of measurement error in ownership. On the other hand, misclassification error is unlikely to be a problem in our case of official firm reports to the Tax Authority on the firm's ownership – a clear, measurable concept reported by professional accountants. This contrasts with the standard cases studied by economists of changes in industry of employment, union membership, or labor force status. In these cases, switching is usually measured in a household survey context by differing answers over time from (potentially different) family members who happen to be home and who are asked questions about one family member's job search, availability, union status, and other employment-related activities.

[20] Kézdi (2003) contains a detailed analysis of autocorrelation and the robust cluster estimator in panel data models.

estimates is somewhat surprising given that the worker characteristics are highly correlated with both wages and ownership, as we documented in the previous section.[21]

Adding firm-specific intercepts, however, greatly diminishes the magnitude of both coefficients, while hardly affecting the estimated wage structure by worker characteristics. The domestic coefficient estimate is -0.096 and the foreign is 0.054. Further adding firm-specific trends reduces the foreign coefficient to close to zero, with a negative sign. The domestic coefficient estimate is still negative, with a magnitude of -0.074, statistically significantly different from zero. Both coefficients in the FE&FT specification have smaller standard errors than in the other specifications, so the issue is not one of precision. But the estimates are not at all robust to these controls for selection bias into ownership type.[22] The hypothesis that the domestic and foreign effects are equal is rejected in all specifications, but the difference between the estimated coefficients is 0.15 under FE and only 0.07 under FE&FT.

Table 7 provides additional estimates that include controls for occupational group of the worker. The estimated coefficients on worker characteristics are somewhat affected by these variables, but they matter little for the estimated impact of private domestic and foreign ownership. At the same time, the ownership coefficients are highly sensitive to the controls for selection bias, but the worker characteristic coefficients are not. The wage structure by worker characteristics that we described in the previous section appears not to result from systematic sorting of workers across firms that pay different wage levels, because any time-invariant firm heterogeneity in wage levels is controlled for in the FE specification while any time-trending heterogeneity across firms is controlled for in the FE&FT.

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[21] These results are little changed by adding interactions between education categories and experience, by estimating separately by gender, or by employing a number of other alternative approaches to estimating earnings functions.

An important and somewhat neglected issue in analyzing the relationship between worker wages and firm characteristics such as ownership is the question of the appropriate unit of observation: the worker or the firm. Analyzing workers exploits the variation in wages among workers and allows their characteristics to be controlled for, so that the composition of employment is held constant. Analyzing firms seems appropriate because ownership is an attribute of the firm and it may be advantageous if the firm-level wage is better measured than are wages at the individual level.[23] Table 8 presents a comparison of some alternative approaches along a number of dimensions: unit of observation (firm or worker), source of dependent variable (firm reports to the Tax Authority, average firm wage constructed from worker data, and individual worker data), sample (manufacturing, "old," or full), and weights on workers when constructing firm-level average wages. The "old" sector includes only firms existing before 1990, and therefore subject to privatization. The first column is the same as Brown et al. (2005), which finds a positive wage impact of privatization to foreign investors and a negligible (at worst slightly negative) impact of domestic privatization. The last column in the Table reproduces our results from Table 6, for comparison purposes. The columns in-between show the results of various changes in the specification and sample. Depending on the choice of specification, one may conclude either that domestic privatization has a large negative impact or only negligible one and that foreign privatization has a large positive impact or only a negligible one, but which specification among these should be preferred is a difficult question.[24]

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[22] We also estimated specifications including other firm-level characteristics, including size, capital-labor ratio, and productivity, but including fixed effects and firm-specific trends rendered these variables statistically insignificant, and they did not affect the coefficients on the ownership variables.

[23] As a check on the possibility that small samples of workers within firms bias our findings, we also estimated all our equations with the sample restricted to employees at firms with at least 100 employees, but the ownership coefficients were qualitatively very similar to those we report for the sample with at least 20 employees.

[24] A similar issue about the appropriate level of observation arises in research on union wage differentials, as discussed by Pencavel (1991), who notes that the few establishment-level studies tend to find lower differentials than those based on individual data. See also DiNardo and Lee (2004), who find no union wage differential using

Because the FE and FE&FT specifications produce such different results from the OLS, it is useful to carry out some specification tests. First, we assess the joint statistical significance of the fixed effects, and then, conditional on including the fixed effects, of the firm-specific trends. The F tests in each case reject the exclusion of the FE and the FT at significance levels of 0.0001. Next, we carry out Hausman tests of the vector of coefficients of the FE model relative to the OLS, and of the FE&FT relative to the FE. Again, these χ^2 tests reject the restricted model in each case. Finally, we carried out pre-ownership-change tests, similar to the "pre-program tests" proposed by Heckman and Hotz (1989) for evaluating alternative non-experimental estimators of the effects of labor market training programs. We implement these tests by adding four variables to equation (1): the first measures whether the firm will become domestic- or foreign-owned 3 years later, the second measures the same event at a 2-year lead, and the third and fourth do the same for a 1-year lead and the year of ownership change. Each of these variables is represented with the notation $\delta_{-\tau i}$ where τ refers to years before the event and i refers to ownership type (d for domestic and f for foreign), so that $\delta_{-2\mathrm{d}}$ is an indicator for 2 years prior to a firm being acquired by domestic private ownership. Our pre-change test consists of simple t tests on each of the first two of these variables ($\delta_{-3\mathrm{d}}$ and $\delta_{-2\mathrm{d}}$ for domestic, and $\delta_{-3\mathrm{f}}$ and $\delta_{-2\mathrm{f}}$ for foreign), as well as an F test on their joint significance.

The results of these tests for the OLS, FE, and FE&FT models are shown in Table 9. The only estimated coefficients that are ever statistically significantly different from zero are for the OLS specification, and the data clearly reject the large foreign effect estimated by OLS. Otherwise, the pre-change test is ambivalent about specifications. In all other cases, however, the estimated ownership effects tend to be small.

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firm-level data on union elections. Although there has been much more research on union than ownership wage differentials, apparently no study of unions uses linked employer-employee data to investigate such differences.

Finally, we turn to the possible impacts of ownership type on skill differentials. Table 10 contains estimates permitting the coefficients on worker characteristics in Table 6 to vary across ownership types. For convenience of interpretation, all these variables are deviated from their sample means. Domestic private ownership appears little different from state ownership in the structure by educational category, but foreign firms appear to offer relatively high rewards to high school and particularly to university graduates. This pattern is attenuated by the addition of firm-specific intercepts and trends, but it remains substantial even in the FE&FT specification. The experience-wage profiles implied by these estimates tend to be flatter (less concave) in domestic private firms and have smaller slopes overall in foreign firms. The implied gender gap is smaller in both types of private firms.

Table 11 contains similar results where occupational categories are included, as in Table 7, but now their coefficients are permitted to vary with ownership type. With unskilled manual worker wages as the base category, the premium for skilled manuals is negative in both types of private firms, in all 3 specifications. While the OLS results imply larger premia for both professionals and managers in foreign-owned firms, the coefficient on professional become tiny and insignificant when firm FE are added, leaving only the managerial group with a substantial premium in foreign firms compared with both types of domestic companies. Domestic private firms offer significantly smaller premia to professionals and associate professionals, according to the FE and FE&FT specifications.

5. Conclusion

The level and structure of wages in Hungary have displayed large changes over the last two decades. After an initial decline, the average real wage has grown steadily and strongly since the mid-1990s, and the wage differentials associated with education and occupation have greatly expanded. Simultaneously, the ownership structure has evolved from completely state-owned to almost completely private, with a very large foreign ownership component. Motivated both by a desire to understand the peculiarities of the Hungarian transition economy and by the large existing literature on differences in wage levels and structure across firms of different ownership types, we have investigated the relationship between wages and ownership using linked employer-employee data for old firms in the manufacturing and services sectors from 1986 to 2003. The employee-side of the data enables us to measure individual worker wages (rather than rely on a firm-level average as in some previous research), to control for individual worker characteristics and changes in the composition of employment that may be correlated with ownership, and to estimate various dimensions of the wage structure, including differentials associated with levels of schooling. The employer-side of the data allows us to measure ownership reliably, and the longitudinal links of employers facilitate some controls for selection bias into ownership type.

We find that simple OLS models imply substantial ownership effects in our data: an approximately 0.18 premium for working in a foreign-owned firm compared to a domestic private company, and a 0.26 premium for state enterprise employees versus those under domestic private ownership. They also imply much higher skill differentials in foreign-owned firms, in particular higher differentials for university education (by 0.27), professional employees (by 0.08) and managers (0.16), relative to those under state ownership. These results control for other worker characteristics, including gender and experience, and for year effects, but they assume no selection bias into ownership types, consistent with much of the literature.

We also estimate models that control for selection based on unobserved heterogeneity through firm fixed effects and firm-specific trend growth in wages. The latter specifications (usually referred to as "random trend models") permit not only idiosyncratic wages at each firm (as in the fixed effects model) but also they allow wages to evolve independently at each firm in a way that is correlated with ownership and with worker characteristics. For example, they permit compensating differentials due to fringe benefits or other work conditions that account for wage differentials to vary across firms not only as a fixed fraction of total compensation, but also to evolve over time according to an idiosyncratic trend for each firm separately.

We find that inclusion of firm fixed effects substantially reduces the differentials in wage levels implied by the OLS estimates and that inclusion of firm-specific trends eliminates the foreign differential entirely. The wage premium of foreign-owned firms relative to domestic private companies falls to about 0.07, and the premium relative to state ownership falls to zero. Differences across ownership types in skill differentials are also reduced in these estimates, but the university and managerial premia in foreign companies remain sizable and statistically significant.

The large variation we receive in results for different specifications motivate us to carry out several specification tests. F tests on the firm fixed effects and firm-specific trends are always highly significant, and Hausman tests imply that the more parsimonious models are incorrectly specified. These results imply that the fixed effects specification is strongly preferred to the OLS, and the specification with trends to the fixed effects specification without trends. We also carry out a pre-program test for each specification to assess whether the specification implies substantial pre-privatization differences in wages, which would suggest residual selection bias. The results from these tests, which we carry out as t tests and F tests on the effects of ownership change 2 and 3 years before the event occurs, are more ambiguous, providing a strong rejection only for the foreign result under OLS. All other test statistics are quite small and statistically insignificant. On the other hand, the only very large wage differential by ownership that we found was the foreign differential under OLS, so the test confirms that this result more likely reflects selection bias than a causal effect.

Taken together, these results imply little or no differences across ownership types in wage levels and structure. There may be a small advantage for employees of state-owned enterprises relative to domestic private firms, but, contrary to the beliefs of many observers and the implications of simple comparisons of mean earnings, neither foreign nor domestic private firms pay a wage premium or provide larger skill differentials, once selection bias has been controlled. One interpretation of this finding is that firm ownership has no causal effect on wage levels and structure, which have evolved in Hungary for other reasons, including changes in product market competition and liberalization of wage determination. An alternative interpretation is that ownership has actually affected wages, but labor market competition has forced state firms to adjust, eliminating any differentials. In this case, we might expect to see some initial ownership differentials that disappear over time as the state closes the gaps. And we might expect that the speed of the state firm's adjustment could be a function of the exposure to private or foreign competition in the particular labor markets where the state firm operates. The first of these expectations may be investigated by estimating the dynamics of the ownership effects for the first few years after ownership change, while the second would require an analysis of spillovers in particular labor markets, for instance defined by region and skill categories. Both of these are potentially fruitful areas of inquiry, but we leave them for future research.

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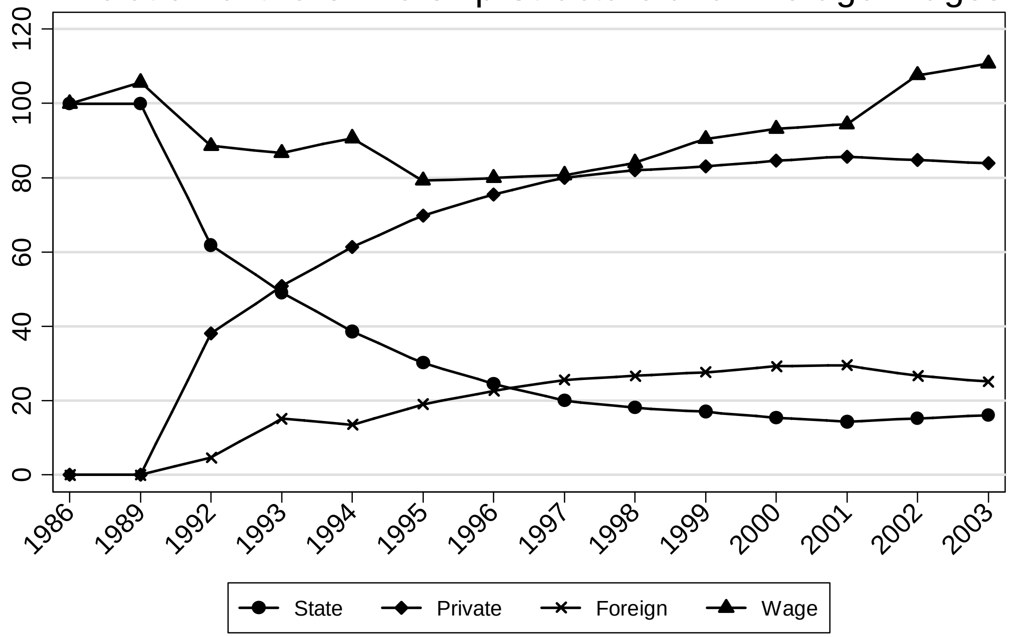
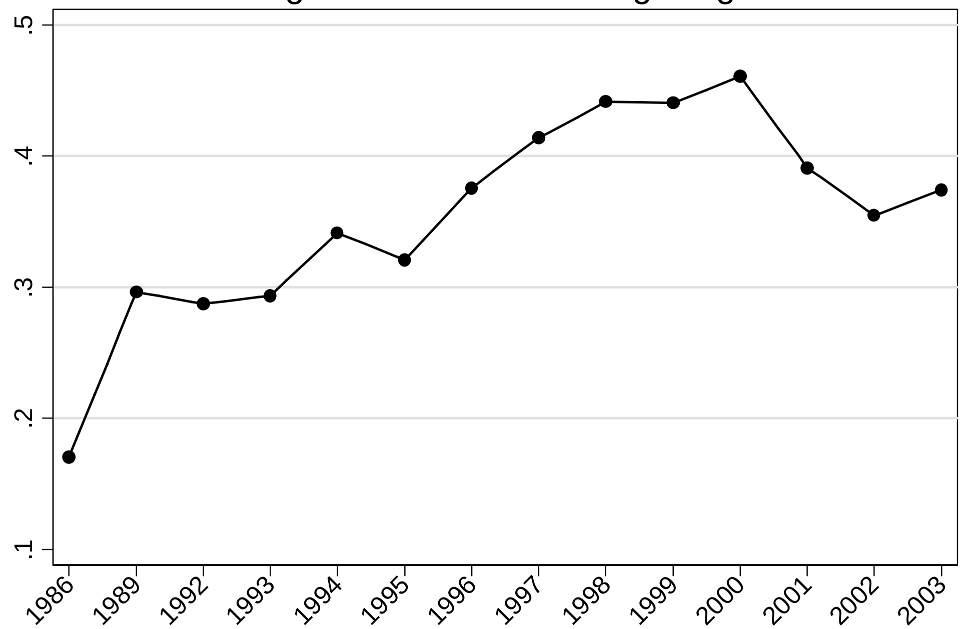
Figure 1:
Evolution of the Ownership Structure and Average Wages



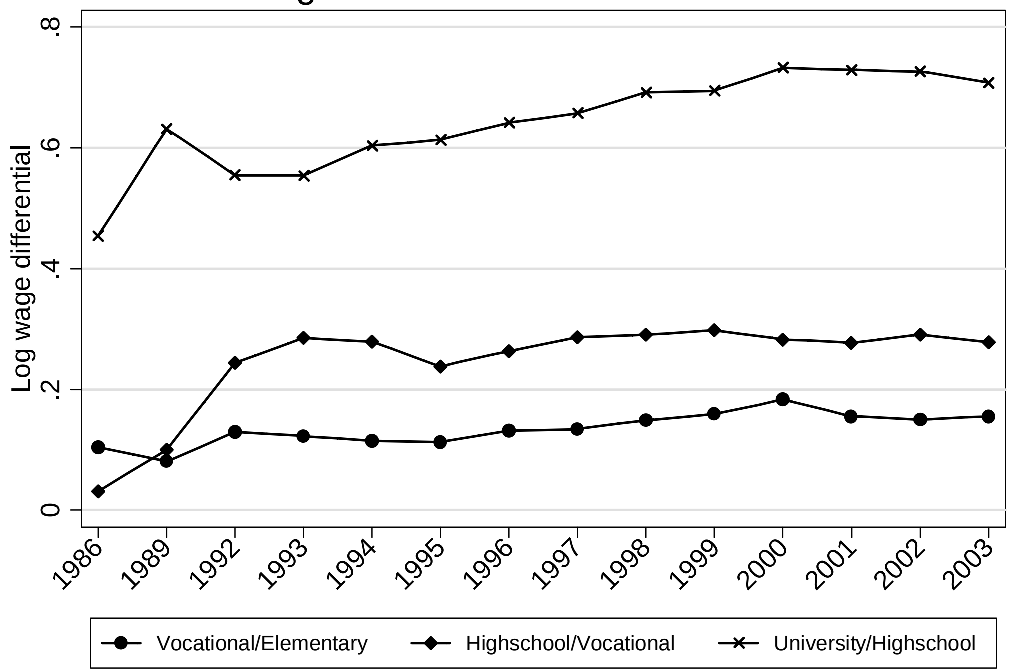
Notes: Number of observations = 1,390,833. State = percent of employees of firms majority state owned. Private = percent of employees of firms majority private. Foreign = percent of firms majority private where foreign is the largest private owner type. The evolution of the average real wage is presented as estimated year effects from a regression including firm fixed effects to control for sample changes (dependent variable = log real wage, normalized at 100 in 1986).

Figure 2: Variance of Log Wage

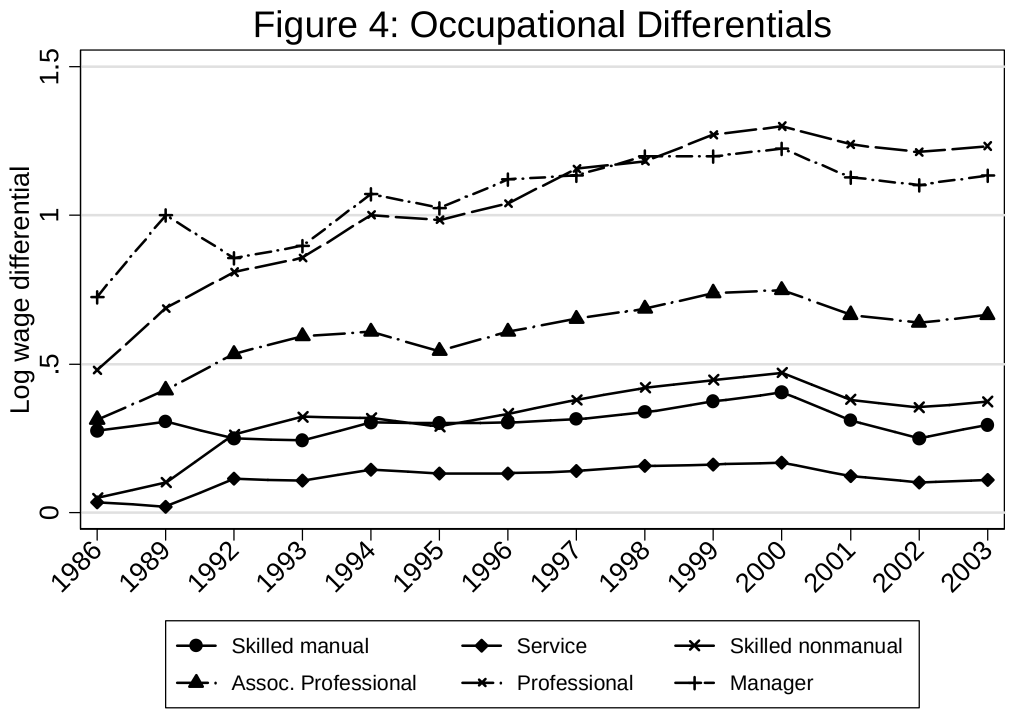


Note: Number of observations: 1,390,833.

Figure 3: Educational Differentials



Notes: Number of observations: 1,390,833. The graphs represent university/high school, high school/vocational and vocational/elementary log real gross wage differentials.



Notes: Number of observations: 1,390,833. The graphs represent log real gross wage differentials of different occupational categories compared to unskilled.

Table 1: Sample Size by Year

Year	Number of Workers	Number of Firms	Total Employment
1986	100,888	3,237	2637.1
1989	107,885	4,031	2272.7
1992	72,993	5,305	1305.7
1993	78,194	5,950	1080.1
1994	100,467	7,672	1417.2
1995	103,815	7,968	1372.5
1996	101,036	7,851	1282.2
1997	87,730	7,440	1230.7
1998	99,626	7,560	1288.1
1999	101,774	8,297	1229.2
2000	112,455	9,449	1266.6
2001	10,865	9,458	1229.7
2002	107,288	5,848	1056.9
2003	105,817	5,266	972.7

Notes: Number of workers = number of workers in sample with information on education, experience, and gender. Number of firms = number of firms with information on ownership and with at least one worker in the given year with information on education, experience, and gender. Total employment = total employment of firms in the sample in thousands (i.e., including nonsampled workers).

Table 2: Number of Ownership Switches

Type of Ownership Change	Number of Firms
State – Domestic	4,437
State – Foreign	293
Domestic – Foreign	383
Foreign – Domestic	360
Always State	3,280
Always Domestic	10,686
Always Foreign	1,835

Notes: Number of firms = 21,381. State = 1 if the firm is at least 50 percent owned by the state in *t*-1. Domestic = 1 if the firm is majority private and domestic owner shareholding is larger than foreign in *t*-1. Foreign = 1 if the firm is majority private and foreign owner shareholding is larger than domestic in *t*-1.

Table 3: Decomposition of Variance of Ln(wage)

	Total	State			Domestic Private			Foreign Private		
	σ^2	S_i	σ_i^2	$(w_i-w)^2$	S_i	σ_i^2	$(w_i-w)^2$	S_i	σ_i^2	$(w_i-w)^2$
1992	0.266	0.626	0.240	0.005	0.330	0.264	0.027	0.044	0.316	0.057
1993	0.301	0.590	0.272	0.005	0.296	0.259	0.056	0.115	0.360	0.043
1994	0.337	0.516	0.283	0.010	0.359	0.335	0.028	0.125	0.366	0.043
1995	0.321	0.411	0.260	0.011	0.423	0.310	0.024	0.166	0.358	0.045
1996	0.370	0.324	0.284	0.013	0.464	0.356	0.023	0.212	0.397	0.054
1997	0.419	0.277	0.324	0.019	0.467	0.376	0.037	0.256	0.423	0.085
1998	0.426	0.225	0.267	0.005	0.482	0.389	0.036	0.293	0.430	0.109
1999	0.451	0.212	0.271	0.011	0.478	0.405	0.041	0.310	0.447	0.120
2000	0.468	0.203	0.296	0.011	0.473	0.429	0.043	0.324	0.440	0.106
2001	0.400	0.198	0.256	0.004	0.473	0.349	0.031	0.329	0.418	0.085
2002	0.379	0.221	0.244	0.005	0.451	0.312	0.026	0.328	0.447	0.097
2003	0.385	0.227	0.248	0.007	0.442	0.311	0.029	0.331	0.462	0.108

Note: S_i is the share of employment (number of observations) of ownership type i (state, domestic private, foreign), σ_i^2 is the variance of ln(wage) for type i, w_i is the mean ln(wage) of type i, and w is the mean ln(wage) in the entire sample.

Table 4: Characteristics of Workers in the Sample, 1992 and 2003

	State		Domestic		Foreign	
	1992	2003	1992	2003	1992	2003
Real Wage	107.3	135.0	82.1	111.0	134.6	197.2
	(70.2)	(109.8)	(57.4)	(113.1)	(107.8)	(214.1)
Female	41.2	38.4	41.1	39.5	46.0	48.1
Education						
Elementary	28.8	19.0	32.8	18.5	29.3	14.0
Vocational	27.5	26.9	35.2	37.7	30.9	27.6
High school	34.2	44.2	24.7	32.4	27.3	37.5
University	9.5	10.0	7.4	11.4	12.6	20.9
Potential experience	22.2	26.3	22.5	25.4	20.8	21.2
	(10.4)	(10.7)	(10.3)	(11.7)	(10.5)	(11.4)
Occupation						
Managers	6.4	11.5	9.7	10.8	6.9	9.5
Professionals	8.6	3.8	6.5	4.1	9.4	9.4
Assoc. professionals	19.0	22.4	11.7	13.9	13.5	22.8
Skilled non-manual	8.7	8.3	10.0	7.5	7.5	7.3
Service	9.6	13.9	6.9	9.3	2.5	5.1
Skilled manual	38.6	32.6	45.3	43.8	49.9	40.7
Unskilled	9.1	7.4	9.9	10.6	10.2	5.2
Observations	38,146	17,060	18,321	60,759	2,769	26,137

Notes: Real wage measured in thousands of 2003 HUF, deflated by CPI. Standard deviations in parentheses.

Table 5: Average Real Wages by Ownership Type and Education

	State		Domestic		Foreign	
	1992	2003	1992	2003	1992	2003
Elementary	79.3	94.9	63.8	75.6	87.6	97.9
	(34.4)	(54.3)	(32.0)	(34.4)	(36.3)	(42.5)
Vocational	92.5	108.8	71.9	83.2	109.6	120.7
	(41.7)	(41.9)	(35.1)	(40.6)	(51.9)	(61.5)
High school	115.3	133.4	94.7	118.0	143.5	178.3
	(57.7)	(76.1)	(60.9)	(91.3)	(77.4)	(142.1)
University	206.1	289.5	170.4	239.8	285.9	398.0
	(135.1)	(237.0)	(106.9)	(244.3)	(200.3)	(348.4)
Observations	38,146	17,060	18,321	60,759	2,769	26,137

Notes: Real wage (deflated by CPI) measured in thousands of 2003 HUF. Standard deviations in parentheses. State = 1 if a majority of the firm's shares are owned by the state. Domestic = 1 if the firm is majority private and domestic owner shareholding is larger than foreign in *t*-1. Foreign = 1 if the firm is majority private and foreign owner shareholding is larger than domestic in *t*-1.

Table 6: Estimated Impacts of Domestic and Foreign Ownership

	OLS	OLS	FE	FE & FT
Domestic	-0.260**	-0.224**	-0.096**	-0.074**
	(0.025)	(0.018)	(0.021)	(0.015)
Foreign	0.176**	0.195**	0.054**	-0.014
	(0.027)	(0.018)	(0.022)	(0.018)
Vocational	-	0.119**	0.130**	0.135**
		(0.005)	(0.003)	(0.004)
High school	-	0.390**	0.328**	0.338**
		(0.010)	(0.005)	(0.006)
University	-	0.977**	0.857**	0.879**
		(0.016)	(0.010)	(0.011)
Experience	-	0.027**	0.027**	0.027**
		(0.001)	(0.000)	(0.000)
$\text{Experience}^2 * 100$	-	-0.039**	-0.038**	-0.038**
		(0.001)	(0.001)	(0.001)
Female	-	-0.200**	-0.200**	-0.197**
		(0.007)	(0.004)	(0.005)
Firm-specific intercepts	No	No	Yes	Yes
Firm-specific trends	No	No	No	Yes
Pr ($\delta_d = \delta_f$)	0.000	0.000	0.000	0.000
R^2	0.141	0.414	0.363	0.357

Notes: Observations = 1,379,453. Dependent variable = ln(real gross wage). *Domestic* = 1 if the firm is majority private and domestic shareholding is larger than foreign in *t*-1. *Foreign* = 1 if the firm is majority private and foreign shareholding are larger than domestic in *t*-1. Elementary is the omitted educational category. FE=specification including firm fixed effects; FT= all variables have been detrended using individual firm trends. All equations include year and region effects. Standard errors (corrected for firm clustering and for loss of degrees of freedom when detrending) are shown in parentheses. R^2: overall for OLS, within for FE and FE&FT. The P values for the F test on the difference between the *Foreign* and *Domestic* coefficients are also reported. ** = significant at 0.01; * = significant at 0.05.

Table 7: Estimated Impacts of Domestic and Foreign Ownership, with Controls for Occupation

	OLS	FE	FE & FT
Domestic	-0.232**	-0.091**	-0.073**
	(0.017)	(0.017)	(0.015)
Foreign	0.184**	0.058**	-0.014
	(0.017)	(0.020)	(0.017)
Vocational	0.064**	0.080**	0.084**
	(0.005)	(0.003)	(0.004)
High school	0.214**	0.158**	0.164**
	(0.008)	(0.004)	(0.005)
University	0.626**	0.504**	0.517**
	(0.013)	(0.008)	(0.009)
Experience	0.022**	0.022**	0.022**
	(0.001)	(0.001)	(0.001)
$\text{Experience}^2 * 100$	-0.032**	-0.032**	-0.032**
	(0.001)	(0.001)	(0.001)
Female	-0.185**	-0.171**	-0.168**
	(0.007)	(0.003)	(0.004)
Skilled manual	0.227**	0.203**	0.206**
	(0.008)	(0.007)	(0.008)
Service	0.074**	0.114**	0.117**
	(0.021)	(0.018)	(0.020)
Skilled non-manual	0.240**	0.208**	0.214**
	(0.011)	(0.008)	(0.010)
Assoc. professional	0.361**	0.309**	0.315**
	(0.015)	(0.011)	(0.012)
Professional	0.453**	0.401**	0.405**
	(0.011)	(0.008)	(0.008)
Manager	0.669**	0.695**	0.710**
	(0.010)	(0.009)	(0.011)
Firm-specific intercepts	No	Yes	Yes
Firm-specific trends	No	No	Yes
Pr ($\delta_d = \delta_f$)	0.000	0.000	0.000
R^2	0.465	0.500	0.451

Notes: Observations = 1,379,453. The specifications are the same as in Table 6 except for the addition of occupational categories. Unskilled manual is the omitted occupation. All equations include year and region effects. Standard errors (corrected for firm clustering and for loss of degrees of freedom when detrending) are shown in parentheses. R^2: overall for OLS, within for FE and FE&FT. The P values for the F test on the difference between the *Foreign* and *Domestic* coefficients are also reported. ** = significant at 0.01; * = significant at 0.05.

Table 8: Firm- versus Worker-Level Estimates

	Estimation Method	**Dependent Variable**					
		Average Wage from Firm Reports	**Average Wage from Worker Survey**		**Individual Worker Wage from Worker Survey**		
Sample		Old Manuf.	Old Manuf. Weighted	Old Manuf. Unweighted	Old Manuf.	All Manuf.	Full
Domestic	OLS	-0.035	-0.199*	-0.171**	-0.103	-0.146**	-0.224**
		(0.020)	(0.082)	(0.018)	(0.056)	(0.053)	(0.018)
	FE	-0.027	-0.060**	-0.051**	-0.050**	-0.040**	-0.096**
		(0.015)	(0.018)	(0.013)	(0.016)	(0.015)	(0.021)
	FE&FT	-0.045**	-0.036	-0.018	-0.027	-0.027	-0.074**
		(0.016)	(0.020)	(0.014)	(0.015)	(0.015)	(0.015)
Foreign	OLS	0.481**	0.247**	0.298**	0.237**	0.211**	0.195**
		(0.036)	(0.058)	(0.027)	(0.046)	(0.046)	(0.018)
	FE	0.307**	0.090**	0.083**	0.060*	0.052*	0.054**
		(0.033)	(0.021)	(0.020)	(0.024)	(0.020)	(0.022)
	FE&FT	0.066*	0.011	0.023	-0.003	-0.006	-0.014
		(0.033)	(0.039)	(0.021)	(0.029)	(0.026)	(0.018)

Notes: These are regression coefficients (standard errors clustered on firms) for alternative specifications in which the unit of observation is the firm or worker, the log wage dependent variable is taken from firm financial reports or the worker survey, the sample varies from old (pre-transition) manufacturing to all nonpublic sectors, and the methods of estimation are OLS, FE (firm fixed effects), and FE&FT (firm-specific intercepts and trends). Average wage from firm data = wage bill/number of employees. Manufacturing = NACE 15-36. Weights are numbers of worker-observations per firm in the Worker Survey. The sample is always restricted to firms with at least 20 employees.

Table 9: Pre-Change Tests

	OLS	FE	FE & FT
Estimated effects 2 and 3 years before privatization			
δ_{-2d}	-0.007	-0.015	-0.011
	(0.049)	(0.022)	(0.014)
t statistic	-0.15	-0.66	-0.80
δ_{-2f}	0.206**	-0.032	0.000
	(0.029)	(0.018)	(0.028)
t statistic	7.24	-1.83	0.00
δ_{-3d}	-0.051*	-0.019	-0.008
	(0.025)	(0.016)	(0.014)
t statistic	-2.03	-1.16	-0.58
δ_{-3f}	0.205**	-0.023	0.004
	(0.031)	(0.022)	(0.026)
t statistic	6.70	-1.03	-0.15
F-Statistics			
Domestic	2.14	0.67	0.35
Foreign	28.70**	1.87	0.04

Notes: The first panel shows the value of the coefficients, standard error (in parentheses) and the value of the *t* statistic for two hypotheses corresponding to tests of the estimated pre-privatization impact of privatization for domestic and foreign ownership, separately: $\delta_{-3d} = 0$, $\delta_{-2d} = 0$, $\delta_{-3f} = 0$, and $\delta_{-2f} = 0$. The second panel shows F-Statistics (P–Values) for two hypotheses corresponding to tests of the estimated pre-privatization impact of privatization for domestic and foreign ownership, separately: $\delta_{-3d} = \delta_{-2d} = 0$, and $\delta_{-3f} = \delta_{-2f} = 0$.

Table 10: Estimated Impacts of Ownership on Wage Structure

	OLS	FE	FE & FT
Base (Omitted category = State)			
Domestic	-0.220**	-0.094**	-0.075**
	(0.017)	(0.020)	(0.015)
Foreign	0.181**	0.039	-0.023
	(0.017)	(0.020)	(0.018)
Vocational	0.142**	0.135**	0.138**
	(0.009)	(0.006)	(0.008)
High school	0.362**	0.321**	0.336**
	(0.013)	(0.011)	(0.012)
University	0.899**	0.835**	0.866**
	(0.027)	(0.019)	(0.025)
Experience	0.034**	0.034**	0.033**
	(0.001)	(0.001)	(0.001)
Experience^2*100	-0.049**	-0.048**	-0.047**
	(0.001)	(0.001)	(0.002)
Female	-0.235**	-0.229**	-0.216**
	(0.010)	(0.011)	(0.012)
Interactions with Domestic Private			
Vocational	-0.037**	-0.004	-0.003
	(0.010)	(0.007)	(0.008)
High school	0.037**	0.016	0.005
	(0.014)	(0.011)	(0.013)
University	0.039	-0.021	-0.032
	(0.031)	(0.021)	(0.027)
Experience	-0.011**	-0.011**	-0.010**
	(0.001)	(0.001)	(0.001)
Experience^2* 100	0.017**	0.016**	0. 016**
	(0.002)	(0.001)	(0.001)
Female	0.084**	0.060**	0.037**
	(0.012)	(0.011)	(0.012)
Interactions with Foreign Ownership			
Vocational	0.017	0.015**	0.013
	(0.013)	(0.009)	(0.010)
High school	0.104**	0.048**	0.029*
	(0.019)	(0.014)	(0.014)
University	0.271**	0.149**	0.119**
	(0.032)	(0.026)	(0.029)
Experience	-0.005**	-0.007**	-0.006**
	(0.001)	(0.001)	(0.001)
Experience^2* 100	0.001	0.006*	0.006
	(0.002)	(0.002)	(0.002)
Female	0.012	0.041**	0.026*
	(0.013)	(0.013)	(0.013)
R^2	0.420	0.368	0.364

Notes: See Table 6.

Table 11: Estimated Impacts of Ownership on Occupational Wage Structure

	OLS	FE	FE & FT
Base (Omitted category = State)			
Domestic	-0.227**	-0.089**	-0.074**
	(0.015)	(0.018)	(0.015)
Foreign	0.172**	0.042*	-0.024
	(0.016)	(0.018)	(0.017)
Skilled manual	0.250**	0.245**	0.250**
	(0.011)	(0.010)	(0.013)
Service	0.137**	0.124**	0.137**
	(0.033)	(0.036)	(0.041)
Skilled non-manual	0.243**	0.233**	0.248**
	(0.015)	(0.013)	(0.015)
Assoc. professional	0.340**	0.332**	0.353**
	(0.020)	(0.019)	(0.021)
Professional	0.433**	0.428**	0.448**
	(0.012)	(0.009)	(0.010)
Manager	0.646**	0.684**	0.698**
	(0.017)	(0.019)	(0.023)
Interactions with Domestic Private			
Skilled manual	-0.027	-0.073**	-0.071**
	(0.014)	(0.011)	(0.014)
Service	-0.125**	-0.013	-0.020
	(0.034)	(0.036)	(0.041)
Skilled non-manual	-0.022	-0.048**	-0.060**
	(0.019)	(0.015)	(0.016)
Assoc. professional	0.047*	-0.053**	-0.069**
	(0.023)	(0.020)	(0.022)
Professional	0.020	-0.077**	-0.101**
	(0.020)	(0.012)	(0.013)
Manager	0.000	-0.027	-0.023
	(0.019)	(0.020)	(0.024)
Interactions with Foreign Ownership			
Skilled manual	-0.048*	-0.079**	-0.084**
	(0.022)	(0.015)	(0.017)
Service	-0.120*	-0.039	-0.061
	(0.049)	(0.044)	(0.048)
Skilled non-manual	0.006	-0.026	-0.048*
	(0.031)	(0.019)	(0.022)
Assoc. professional	0.015	-0.020	-0.055*
	(0.037)	(0.024)	(0.027)
Professional	0.076*	0.009	-0.031
	(0.033)	(0.022)	(0.021)
Manager	0.160**	0.147**	0.133**
	(0.033)	(0.026)	(0.029)
R^2	0.474	0.458	0.458

Notes: See Table 6. All equations include year and region effects and all other variables from Table 10.