

Foreign Ownership and the Distribution of Wages in Hungary, 1992-2000: An Unconditional Quantile Decomposition Approach¹

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Abstract

With the help of a rich linked dataset on both firms and workers of the Hungarian corporate sector, this paper analyzes how changes in foreign direct investment contributed to changes in the unconditional wage distribution at different quantiles between 1992 and 2000. After transition, Hungary experienced an extraordinary amount of continuous FDI inflow during the nineties, while earnings inequality increased by close to seventy percent in just ten years, compared to its 1989 level. The role of FDI in inequality changes is partialled out by a detailed decomposition of log wage changes based on a recently developed method by Firpo et al. (2009) that extends the standard Oaxaca-Blinder decomposition to unconditional quantiles of the distribution. I find that at every point in time, the share of employees of foreign-owned firms has a positive and significant wage level effect at every unconditional quantile, and these effects are inequality enhancing for men while they have an ambiguous effect on the unconditional dispersion for women. FDI contributed strongly to wage changes at every part of the distribution through an increased foreign employment share in the economy, but not through changes in the returns to being employed by foreign-owned firms. However, it played only a moderate role in the growth of inequality.

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1. Introduction

Wage inequality has been one of the most studied topics in all of economic research during the last two decades. Bound and Johnson (1992), Katz and Murphy (1992), Juhn et al. (1993), Lemieux (2006) and Autor et al. (2008) are among the seminal contributions documenting the rise in inequality and some leading explanations behind the trends in the United States. Katz and Autor (1999) summarize the early literature on the U.S. and a few other economies, and Lemieux (2008) provides a recent overview. Both of the last two papers point out that the U.S. economy is a bit of an outlier and that there is significant variation across countries in the extent and patterns of the inequality increase. A rather thinner, but significant, strand of literature focused on the wage effects of foreign ownership – in particular, of foreign acquisitions – to find that in most settings foreign-owned employers pay higher wages on average, keeping everything else constant (e.g. Conyon et al., 2002; Heyman et al., 2007; Huttunen, 2007; and Girma and Görg, 2007; among others). This paper is at the intersection of these two areas of labor market research, as it investigates the effect of foreign ownership on the wage distribution as a whole.

Despite vivid interest regarding wage inequality in the U.S., the wage distribution has been carefully studied in rather few countries, with a heavy emphasis on developed economies. In particular, among Central and East European economies, only few thorough studies exist. Keane and Prasad (2006) study the effect of the Polish transition on the structure of earnings, and Ganguli and Terrell (2006) analyze Ukraine. However, the former use data only through 1996, the latter only two cross-sections (1986 and 2003) and both have little information on the employer side. Concerning Hungary, with the exception of the early years of transition (Kertesi and Köllő, 1997), there has been little research on inequality in general, although some aspects of wage differentials have received attention.²

² For example, Jolliffe and Campos (2004) estimate the effects of market liberalization on the gender wage gap; Kertesi and Köllő study skill differentials (2002) and industrial wage differences (2003a, 2003b); Neumann

It would be profitable to aim at a more pronounced presence of Central and Eastern Europe on the map of earnings inequality research, since the structure of wages in the transition and accession economies of the region have changed dramatically in the last two decades 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 human capital show large increases in every country where they have been studied. Following the transition, 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 tended to be compressed under socialism (e.g., Kornai, 1992). At the same time, these countries have experienced massive organizational changes associated with price and trade liberalization, privatization of most types of enterprises, evolution in the institutional environment, and opening to the global economy, particularly to foreign direct investment (FDI).

It is this last factor, the inflow of foreign capital in form of greenfield investments and acquisitions that the research in this paper is focused at. According to the OECD (2000), during the nineties, Hungary received the largest amount of FDI in the region, and the interest of foreign investors has not languished in the subsequent decade either. The period of fast growth in wage inequality coincided with large-scale privatization and a huge inflow of foreign direct investment, and with the arrival of foreign investors new wage-setting strategies appeared in the country. These rapid changes provide a useful context for investigating the following research questions: Did the ever-increasing inflow of FDI contribute to the dramatic jump in wage inequality? If it did, which parts of the distribution

(2002) explores the effect of collective wage bargaining; Köllő and Nagy (1996) study the effects of unemployment on earnings; and Earle et al. (2011) investigate the effects of foreign ownership on average wages and the wage structure. The Labor Market Yearbook (2000) contains an overview of the evolution of wages during transition. Preliminary results spanning the entire transition era and the most recent years from an analysis that is concerned with the complete wage distribution show a dramatic increase in the dispersion of earnings, with a rate of growth comparable to U.S. trends (Antal, 2011).

were affected the most heavily? Do we observe a heterogeneous effect similar to the decline in unionization in the U.S., where the impact of changes in the share of union workers was reducing inequality below the median, but widening wage dispersion in the top half of the earnings distribution?³ Is it only the rise in the foreign share in employment that matters – i.e. a composition effect –, or was it coupled with a wage structure effect, that is, with a change over time in the labor market return to being employed by foreign owned companies?

Hungary provides a particularly interesting and fruitful case for this analysis, one with the potential to provide lessons of broader importance to scholars interested in a variety of economies. Unlike many other countries of Central and Eastern Europe, the liberalization and privatization processes were relatively quick, and they were largely completed by the early to mid-1990s (e.g., Frydman et al. 1993; Mihályi, 1997). Both the ownership structure – the predominance of concentrated outside ownership in large corporations – and the openness of the Hungarian economy quickly became much more similar to developed European economies than elsewhere in the formerly socialist bloc (e.g., Brown et al. 2006). The overall business and policy environment also converged more quickly to European norms, while elsewhere in the region problems of corruption and bureaucratic interference in business tended to be more persistent (e.g., Kaufmann et al. 2003). Hungary, besides having an institutional structure that provides a useful setting for the analysis of wage inequality in transition in general, and of the role of FDI in shaping the wage distribution in particular, also offers a unique database containing linked observations on employees and employers covering the time periods before, during, and after the transition, and containing rich information on both workers and their workplaces.

This paper moves beyond most previous studies of the relationship between foreign direct investment and wages, in estimating the direct contribution of changes in both the

³ See for example Firpo et al. (2007).

distribution of, and the returns to foreign ownership to changes in the unconditional distribution of wages with the help of a recently developed method by Firpo et al. (2007, 2009). To estimate this contribution, any method has to capture two effects of FDI that might work at the same time. The first is a between-group effect represented by the difference in mean wages of otherwise comparable employees of foreign and domestic firms. The second is a within-group effect that is generated by potential differences in conditional foreign wage premia at different quantiles of the conditional wage distribution of worker groups defined by individual characteristics other than foreign control. Running usual conditional quantile regressions would only capture this latter effect, while the major part of the literature on FDI and wages focuses on the first one. Firpo et al. (2009) merge the concepts of influence functions and quantile regressions to be able to estimate the partial effect of changes in covariates on the unconditional quantiles of the wage distribution. This regression framework is then extended by Firpo et al. (2007) to decompose changes in the unconditional wage quantiles over time and measure the contributions of single covariates through a composition and a wage structure channel which are both allowed to be heterogeneous across quantiles.

2. Foreign Ownership and Wage Dispersion: Current State of Knowledge

Studies of wage inequality typically exclude peculiarities of employers and concentrate on individual characteristics of the worker, like education, experience, occupation and gender. Usually, region and industry controls are also involved in the analysis; however, these are rather indicators for heterogeneity in labor markets and not for heterogeneity of firms. One notable exception that received a lot of attention in the wage

inequality literature on the U.S., and that might be viewed as both a worker and a firm attribute, is union status.⁴

Including firm characteristics when examining changing patterns of wage dispersion might be fruitful *ex ante* for at least two reasons. First, an emerging strand of literature delivered evidence on the substantial role of between-establishment wage dispersion in shaping the overall distribution of wages.⁵ Second, a large chunk of the rise in wage inequality is still unaccounted for even after controlling for the usual suspects for possible explanations like skill-biased technological change, import competition, worker composition effects, and changes in labor market institutions, such as the minimum wage legislation and the degree of unionization.⁶

Turning to the literature on FDI and wages, the relationship between foreign ownership and average wages has been investigated to a large extent, but the same is not true for other moments of the wage distribution. Authors using firm-level data typically find a positive foreign premium regarding the conditional expectation of wages,⁷ while a smaller fraction of papers based on linked employer-employee data tend to find a smaller or even a zero causal effect after controlling for various individual and firm characteristics, and for unobserved heterogeneity as much as the data permit to do so – with the exception of Hungary, where the effect is large even in the linked sample.⁸ Note that in the case of the

⁴ See DiNardo et al. (1996), Firpo et al. (2007, 2009) and DiNardo and Lemieux (1997). In the paper, I rely mostly on methods that were developed in this series of studies addressing the impact of deunionization on wage inequality.

⁵ Among others, see Davis and Haltiwanger (1991), Abowd et al. (1999), Dunne et al. (2004), Haltiwanger et al. (2007).

⁶ Skill endowments and skill prices, account only for a small fraction of both the level and the change in overall wage inequality. Juhn et al. (1993) find that in the U.S., unobserved skill quantities and prices account for 56% of the total growth in the log 90-10 wage differential of men, between 1964 and 1988. Autor et al. (2005) show that the same measure is 63% in the 1988-2003 period.

⁷ Firm-level studies include Conyon et al. (2002), and Girma and Görg (2007), both on the UK; Aitken et al. (1996) on Mexico, Venezuela, and the US; Feliciano and Lipsey (2006) on the US; Lipsey and Sjöholm (2004) on Indonesia; and Brown et al. (2008) on FDI entry through privatization in Hungary, Romania, Russia, and Ukraine.

⁸ See Martins (2004) and Almeida (2007) for Portugal; Heyman et al. (2006, 2007) for Sweden; Huttunen (2007) for Finland; Andrews et al. (2007) for Germany; Martins and Esteves (2008) for Brazil; and Earle et al. (2011) for Hungary.

first moment, studies concerned with the causal effects of foreign acquisitions and/or divestments on the conditional expectation of wages implicitly estimate an unconditional effect, too. It can be shown that in an OLS regression, the estimated coefficient on the foreign dummy has a dual interpretation: it gives the expected return to a worker employed by a foreign firm relative to a worker with similar characteristics employed by a domestic firm; but it also measures the marginal effect of an increase in the share of foreign employment on the unconditional mean of wages.⁹ Still, this does not tell us anything about a possible change in the shape of the unconditional distribution of wages.

Of course, a difference in average wages at the firm-level between domestic and foreign firms contributes directly to between-firm wage inequality and thus to overall inequality as well, but mostly, the focus of these studies is to disentangle spurious correlation and causal effects, and to measure a "true" foreign wage premium and not to quantify the contribution to inequality. Furthermore, as discussed in Section 1, the difference in conditional first moments is only a between-group effect on the unconditional wage distribution, but there might be another, within-group effect potentially implying different conditional wage premia at different points of the conditional wage distribution. To evaluate the impact of foreign ownership on the unconditional wage dispersion, one has to estimate these two effects simultaneously that can be performed in an unconditional quantile regression framework applied later in the paper.

There are some recent studies that move beyond the relationship of FDI and the conditional grand mean of wages. However, in most cases, the analysis is only extended to wage structure effects in that the work force is divided into a few skill groups, and the authors estimate foreign wage premia separately for each of these. For example, Huttunen (2007) examines how the foreign acquisition wage effect varies by educational groups of Finnish

⁹ This is only true if the foreign dummy enters the regression independently – i.e. not interacted with other covariates –, otherwise one has to integrate over the distribution of other covariates to get to the unconditional interpretation.

plants and finds that the magnitude is increasing with the level of schooling. Almeida (2007) shows that after controlling for selection, foreign acquisitions result in only modest wage gains for Portuguese workers, but the difference is increasing in skill. Eriksson and Pytliková (2011), studying a single cross-section of Czech linked employer-employee data, disaggregate workers into white-collar and blue-collar employees and find that FDI benefits both groups, with a higher average wage gain for the former group.¹⁰ These studies all estimate a price effect of FDI on the wage distribution, but since the focus is not on quantifying to exactly what extent changes in the ownership structure contributed to changes in the wage distribution, they ignore the composition effect and the time dimension.

Following a different approach, Eriksson et al. (2009) look at the evolution of within-firm and between-firm inequality in the Czech Republic, and find a quite robust positive effect of foreign ownership on both. However, they only have data for the 1998-2006 period, and thus lose valuable information on early-transition years, where the most important changes affecting the wage structure presumably happened. Also, their dependent variables are the within-firm and between-firm conditional standard deviation of wages, so they do not answer the question how the inflow of foreign capital affected dispersion in the unconditional distribution. As noted earlier, this research contributes to the literature by shedding light on this latter aspect of the relationship between FDI and labor market outcomes.

3. Description of Data

The main body of data used in this paper comes from the Hungarian Wage Survey (HWS), a yearly survey on employees – augmented with some information on the employer – conducted by the Central Statistical Office. To assemble the linked employer-employee

¹⁰ Using Swedish data, Heyman et al. (2006) distinguish three skill groups, and find – somewhat surprisingly – that foreign acquisitions result in wage growth only for the top occupational groups, in particular for CEOs and managers, moreover, it seems to be a consequence of acquisition, and not a genuine foreign effect. In contrast, Earle et al. (2011) estimate the foreign wage structure effect in Hungary, by interacting the foreign acquisition dummy with various individual characteristics and show that every worker group experiences an increase in average wages, with extra premia for the high-educated.

dataset (LEED) necessary for the analysis, the worker-level HWS files are linked with the help of a firm identifier to administrative firm-level data collected by the Hungarian Tax Authority (HTA). The HTA database contains the complete balance sheet and income statement of firms with double-entry book keeping in the Hungarian business sector. The inclusion of employers in the HWS based on their size has changed over the years. All business units were surveyed in 1986, 1989; the size threshold of sample inclusion was at least 20 employees between 1992 and 1995; a random sample of firms with 11-20 employees was added to the group of larger firms between 1996 and 1999, which was extended to the threshold of 5 employees thereafter. Since the foreign share of firms varies by size, and the size distribution of firms is truncated at different points in different years, I only include firms with more than 20 employees in any given year to insure sample consistency. Within business units, workers were sampled representatively, based on a random sampling design.¹¹ The linked data are a panel in firms, but not in workers, although it is possible to follow the majority of individuals that do not change employers over time exploiting the birth-date-based sampling design and the rich set of observed characteristics.

I use two sets of sample weights. The first is a within-establishment and within-occupational-group worker weight to account for the different degree of representation of the blue- and white-collar workers within establishments. The second is a company-level multiplier that weights up the sample to the total employment of the corporate sector of Hungary. The final sample weight is the product of the individual and the firm weights. In addition, whenever it is necessary for constructing counterfactual samples, I weight

¹¹ In 1986 and 1989, all senior managers were included, and a random sample of the rest of the professionals was selected on the basis of the socialist wage grid; the first and every fifth employee in 1986 and every tenth in 1989. In case of manual workers, the survey covered the first and then every seventh worker of each wage group in 1986, while the first and every tenth person in 1989. In 1992 and 1993, every blue-collar worker born on the 5th or 15th of any month and every white-collar worker born on the 5th, 15th or 25th was surveyed. This scheme was maintained after 1993 for firms above a certain size threshold, and all employees' information were required from sampled companies smaller than the threshold. The size limit was 20 employees from 1994 to 2001, and it was raised to 50 employees thereafter

observations by probability weights that rescale the actual distribution of individual characteristics. How these weights are estimated, is described in detail in Section 4. Of course, once applied, probability weights are always multiplied by sample weights.

The HWS contains information on earnings and its various components (e.g., basic wage, bonuses, temporary payments, commissions and allowances), working hours, demographic and human capital variables, four-digit occupation codes and plant-level location information. From the HTA data, I use information on equity share to construct a foreign ownership dummy, and industrial affiliation. The data were cleaned both at the individual- and the firm-level extensively. Variable definitions and classifications are harmonized over time, with special attention to the synchronization of the pre- and post-transition parts of time series. The common issue of spurious firm entry and exit was addressed with the help of another dataset that provides administrative information on boundary changes, entry and exit. Exit and entry rates were improved by detecting more than 3,000 erroneous identifier changes.

The main variable of interest, ownership, was carefully cleaned both at the share level, and after defining ownership dummies based on relative equity shares. At the share level, “roundripper” observations – where shares switch owners for a single year – were recoded, and impossible changes – like the increase in state ownership shares – were cleaned where possible, or were set to missing. I use a simple majority definition of foreign ownership, that is, the firm is considered to be foreign owned if at least fifty percent of its equity is owned by foreign investors. Roundripper observations were also cleaned at the dummy level, and I filled up missing values in the middle of long series of the same owner type. Of the sample of firms with more than 20 employees, I dropped 855 which experience more than two changes in ownership status because in general these tend to have unreliable ownership data.

I excluded part-time employees since they are only observed after 2002, and I only kept individuals with an age of more than 15 and less than 74 years. Two-digit industries that do not have any foreign presence at all are excluded from the analysis (NACE 42 and 91), as well as NACE codes 75, 80 and 85 (public administration). The final sample comprises 2.5 million worker-year observations, and a total of 25,000 unique firms, covering two decades between 1986 and 2008. I present the evolution of inequality for all these years, but I only investigate the effect of foreign ownership for the period between 1992 and 2000, the decade of the largest increase in wage dispersion. Table 1 summarizes by-year information about sample sizes and the degree of representation. On average, the sample includes 100,000 workers per year, representing 1.1 million employees. The third column shows that this is close to seventy percent of total employment by enterprises with more than 20 employees, as measured by the sum of employment of these firms in the HTA dataset. I demonstrate in Figure 1 that the linked sample is also doing a very good job in tracking down patterns of foreign ownership shares in both number of firms and number of employees in the full business sector data. In terms of number of firms, foreign penetration reaches 20 percent by 2008, while the share in total employment is close to 40 percent among companies above the 20-employee limit.

I provide some descriptive statistics of the sample in Appendix Table 1. The share of workers employed by foreign-owned enterprises in 2000 was six times of its value in 1992. Raw average earnings were 31 percent higher in foreign than in domestic firms already in 1992, and that gap increased to 53 percent by 2000. By the end of the nineties, foreign firms employed a younger and higher skilled workforce. The share of women was by close to ten percentage points higher than in domestic firms in the early years of transition and basically did not change in the following decade. Foreign businesses were concentrated heavily in machine and equipment manufacturing – especially in the starting period of the inflow of

foreign capital –, and they employed more skilled manual workers accordingly, but also more professionals. It is important to note that FDI affected all industries, with the exception of mining.

4. Estimation Methods

The ultimate goal of decomposition exercises is to quantify the effect of changes in the composition of observable and unobservable factors, and the effect of changes in labor market returns to these factors (or in other words, changes in the wage structure) on the unconditional distribution of wages. That is, we aim to decompose the overall change in some functional $v(G_w)$ of the unconditional wage distribution characterized by the cumulative distribution function (CDF) $G_w(\cdot)$, into a factor due to underlying changes in the composition of individual characteristics, and into a factor due to changes in the wage structure where the set of individual characteristics might include both observable and unobservable elements, and by wage structure we might think of a process that rewards both types of characteristics. This is called an aggregate decomposition in the literature,¹² and is given by $\Delta_O^{v(G)} \equiv \Delta_C^{v(G)} + \Delta_S^{v(G)}$, where the left-hand side is the overall difference in some distributional attribute of wages between two states (e.g. worker groups or periods of time), and the first term of the sum represents composition effects while the second the wage structure effects.¹³ In most applications, however, we are interested in the contribution of some individual characteristic of interest to both the composition and the wage structure effects – as in this paper, I want to quantify the contribution of foreign ownership –, so we would like to further decompose $\Delta_C^{v(G)}$ and $\Delta_S^{v(G)}$ by means of a detailed decomposition.

¹² For an exhaustive survey of decomposition methods see Fortin et al. (2011). I follow to a great extent their notation and terminology in this paper.

¹³ Consider the case of the well-known mean decomposition proposed in the seminal papers of Oaxaca (1973) and Blinder (1973). In their method, $v(G)$ equals μ , the unconditional mean of wages, and the wage generating process is assumed to be $w = X\beta + \varepsilon$, where X is a vector of observed characteristics and ε is an idiosyncratic error term. Then, after setting some identifying assumptions, the estimated OB decomposition in its simplest form between two groups 1 and 2, is given by $\Delta_O^\mu \equiv \Delta_X^\mu + \Delta_S^\mu = (\bar{X}_1 - \bar{X}_2)\hat{\beta}_1 + \bar{X}_2(\hat{\beta}_1 - \hat{\beta}_2)$, where \bar{X} -s are sample averages and $\hat{\beta}$ -s are estimated by OLS. It is also straightforward to compute the contribution of each element in X to both parts of the decomposition.

To perform decompositions, counterfactual states of the world have to be constructed keeping one or more factors fixed while letting the others change to partial out their effect. The aim of this paper is to partial out the composition and wage structure effects of foreign ownership. The related counterfactual exercises are of the form: How would the distribution of wages look like in some end period if every factor was kept at some base period's level except the proportion of workers employed by foreign-owned companies (or alternatively, the returns to being employed by a foreign-owned company)? The difference between an actual unconditional wage distribution, and the counterfactual wage distribution gives either the composition or the wage structure effect of foreign investment, depending on what dimension of foreign ownership was allowed to change over time.

Once the counterfactual and actual distributions are specified and identifying assumptions are set, one can estimate either the corresponding distributions non-parametrically, or various functionals of the distributions (like quantiles, the variance and other inequality measures) with the help of a parametric model. I will both decompose densities of wages, and quantiles and interquantile ranges of wage distributions to estimate the contribution of foreign ownership to changes in wage inequality. In section 4.1, I briefly summarize the reweighting decomposition method of DiNardo et al. (1996) and of DiNardo and Lemieux (1997) – DFL and DL henceforth, respectively – which I apply to construct counterfactual densities and to perform density decompositions. The DFL method is designed to identify composition effects, while the DL method is an extension to identify wage structure effects. In section 4.2, I describe the procedure recently developed by Firpo et al. (2007) – FFL henceforth – that combines reweighting and recentered influence function (RIF) regressions to decompose quantiles and inequality measures.

4.1 Counterfactual Wage Distributions

As I examine changes in the wage distribution over time, between two periods, let $t = \{b, e\}$ denote time, where b refers to some base period, and e to some end period. Log

wages w are determined at any point in time by the wage structure function, $w_t = s_t(X, \varepsilon)$,¹⁴ where $X = [X_t]_{t=1}^k = [f, Z]$ is a vector of worker and firm characteristics with $X_1 = f$ denoting the foreign ownership dummy, and $Z = [Z_t]_{t=2}^k$ the rest of covariates, such as education, gender, occupation, potential experience, region of workplace and industrial affiliation of the employer; while ε measures unobserved individual heterogeneity.¹⁵ Let $G_{X,t}(X)$ denote the (marginal) CDF of covariates in t , while $G_{w|X,t}(w|X)$ the conditional, and $G_{w,t}(w)$ the unconditional CDF of wages in t . Let $g_t(\cdot)$ refer to the corresponding densities, for example, $g_{w,t}(w) = \frac{dG_{w,t}(w)}{dw}$.

The unconditional density of log wages in time t can be written as

$$(1) \quad g_{w,t}(w) = \int g_{w|X,t}(w|X) dG_{X,t}(X) = \int g_{w|X,t}(w|f, Z) dG_{X,t}(f, Z).$$

Since the joint distribution of covariates, $G_{X,t}(f, Z)$, can be expressed as $G_{f|Z,t}(f|Z)G_{Z,t}(Z)$, equation (1) turns into

$$(2) \quad g_{w,t}(w) = \iint g_{w|X,t}(w|f, Z) dG_{f|Z,t}(f|Z) dG_{Z,t}(Z).$$

One counterfactual distribution of interest is the hypothetical distribution of wages that would prevail if the proportion of workers employed by foreign-owned firms was fixed at its base period level, but other individual attributes, and the conditional density of wages were allowed to change to their end period values. Let the asterisk in the upper index always denote counterfactual states, then we have the following counterfactual density of wages for the composition effect exercise (marked by upper index C):

¹⁴ Since this subsection is about the non-parametric estimation of densities, I will only specify the form of the wage structure function in subsection 4.2.

¹⁵ To simplify the discussion in this section, all characteristics of the employer are considered as individual characteristics of the employee.

$$(3) \quad g_{w,\varepsilon}^{*,c}(w) = \iint g_{w|x,\varepsilon}(w|f, Z) dG_{f|z,b}(f|Z) dG_{z,\varepsilon}(Z).$$

Note that this formulation implicitly involves a very strong assumption called the invariance of conditional distributions by Fortin et al. (2011). By equation (3), we are ruling out self-selection into foreign ownership status based on ε , as well as general equilibrium (or spillover) effects of foreign investment, since these are both assumed away by leaving the conditional distribution of wages unaffected while changing the distribution of foreign ownership.

Following the DFL technique, the conditional base period distribution of foreign ownership can be substituted by properly reweighting its end period conditional distribution to get

$$(4) \quad g_{w,\varepsilon}^{*,c}(w) = \iint g_{w|x,\varepsilon}(w|f, Z) \Psi^c(f, Z) dG_{f|z,\varepsilon}(f|Z) dG_{z,\varepsilon}(Z),$$

where the reweighting function is

$$(5) \quad \Psi^c(f, Z) \equiv \frac{dG_{f|z,b}(f|Z)}{dG_{f|z,\varepsilon}(f|Z)} = f \cdot \frac{Pr_b(f=1|Z)}{Pr_\varepsilon(f=1|Z)} + (1-f) \cdot \frac{Pr_b(f=0|Z)}{Pr_\varepsilon(f=0|Z)}.$$

The main advantage of this formulation is that the conditional probabilities of ownership status $Pr_b(f=1|Z)$ and $Pr_b(f=0|Z)$ can be readily estimated by specifying a probit or logit model for the probability of being employed by a foreign-owned company conditional on Z in both periods. The predicted probabilities are then used to estimate $\Psi^c(f, Z)$ for every observation of the end period.

The contribution of foreign ownership to the composition effect of X on the change in the unconditional density of log wages is then given by the difference between the actual

unconditional density of log wages in the end period, and the counterfactual density defined by equation (4). That is

$$(6) \quad \Delta_{C_f}^{g(w)} \equiv g_{w,s}(w) - g_{w,s}^{*,C}(w).$$

To demonstrate the wage structure effect of a binary variable on the density of wages, DiNardo and Lemieux (1997) follow a different path, but applying a similar reweighting procedure that led to expression (6) for the composition effect. The hypothetical state of the world we would need is one where every worker was paid under the wage structure of domestic companies. The distribution of wages in this state for period t could be constructed as

$$(7) \quad g_{w,t}^{*,S}(w) = \int g_{w|x,t}(w|f=0, Z) dG_{Z,t}(Z).$$

Now since we do not observe the wage structure, $g_{w|x,t}(w|f=0, Z)$, for workers of foreign firms, we cannot estimate the counterfactual density in this form. Thus, DiNardo and Lemieux suggest reweighting the sample of domestic workers so that the distribution of Z in this subsample reflects the distribution in the entire sample. Applying Bayes' Law, we can write

$$(8) \quad G_{Z,t}(Z) = \frac{G_{Z|f,t}(Z|f=0) \cdot Pr_t(f=0)}{Pr_t(f=0|Z)} = G_{Z|f,t}(Z|f=0) \cdot \Psi_t^S(Z),$$

where $\Psi_t^S(Z)$ denotes the reweighting function used for the wage structure effect exercise. It can be estimated the same way as $\Psi^C(f, Z)$ by running a binary outcome regression for $Pr_t(f=0|Z)$, and by replacing $Pr_t(f=0)$ by the proportion of workers in the sample employed by domestic companies. The estimable counterfactual density is given by

$$(9) \quad g_{w,t}^{*,S}(w) = \int g_{w|x,t}(w|f=0, Z) \Psi_t^S(Z) dG_{Z|f,t}(Z|f=0),$$

and it is only estimated for the subsample of domestic workers.

Note that the expression in (9), including the reweighting function, is time dependent, since the counterfactual thought experiment is different by nature from the one for the composition effect. Nevertheless, we can define the wage structure effect of foreign ownership on the density of wages over time by

$$(10) \quad \Delta_{S_f}^{g(w)} \equiv [g_{w,s}(w) - g_{w,s}^{*,S}(w)] - [g_{w,b}(w) - g_{w,b}^{*,S}(w)].$$

The interpretation of definition (10) is that once we know the part of the observed density of wages that can be attributed to the different pay schemes of foreign and domestic employers in both periods, the wage structure effect of foreign ownership on changes in the wage distribution is simply the difference of these two parts.

The only missing element is the estimation of the actual and the counterfactual densities. In the first step, I obtain $\hat{\Psi}^c(f, Z)$ and $\hat{\Psi}_t^S(Z)$ as described above, and in the second step, densities are estimated by a kernel density estimator of the form

$$(9) \quad \hat{g}_{w,t}(w) = \frac{1}{h \sum \omega_t} \sum \omega_t K\left(\frac{w - W_t}{h}\right) \quad \text{and}$$

$$(10) \quad \hat{g}_{w,t}^*(w) = \frac{1}{h \sum \omega_t} \sum \omega_t \hat{\Psi}(\cdot) K\left(\frac{w - W_t}{h}\right),$$

where ω_t are sample weights, h is the bandwidth of the kernel, $K(\cdot)$ is the Epanechnikov kernel function, while $\hat{\Psi}(\cdot)$ denotes either $\hat{\Psi}^c(f, Z)$ or $\hat{\Psi}_t^S(Z)$, depending on which counterfactual density we would like to estimate.

4.2 Decomposition Based on Unconditional Quantile Regressions

Section 4.1 was instructive as (i) it showed how the DFL reweighting approach works in general, (ii) it proposed a tool for visually inspecting the difference between actual and properly defined counterfactual densities to get a hint about the nature of foreign composition and wage structure effects, and (iii) it provides a basis for decomposing any usual distributional statistic or inequality measure. The last point is straightforward, because once we estimated the densities non-parametrically; interquantile differentials, the variance, the Gini coefficient or other measures of dispersion can be computed. However, the method has an important limitation: the generalization to more and to non-binary covariates is cumbersome, especially in case of the wage structure effect.¹⁶ This is why I will work with a hybrid method proposed by Firpo et al. (2007) and Fortin et al. (2011) that combines DFL reweighting and a method called recentered influence function (RIF) regression to perform detailed decompositions similar to a standard OB decomposition on quantiles and inequality measures. The FFL decomposition is easy to implement, has good asymptotic features, is straightforward to interpret,¹⁷ and is computationally feasible for large datasets such as the LEED used in this paper.¹⁸ First, I describe the concept of RIF regressions, as introduced by Firpo et al. (2009), and then the decomposition method that builds on RIF projections and DFL reweighting.

The main goal of detailed quantile decomposition is to quantify the effect of covariates on various parts of the wage distribution. Thus, the RIF idea is based on the concept of the influence function, a tool introduced in the robust estimation literature by Hampel (1974) to measure the effect of small perturbations of an underlying distribution on

¹⁶ See more about the advantages and limitations of the DFL method versus other approaches in Fortin et al. (2011).

¹⁷ Except for wage structure effects that suffer from the same omitted group problem as the OB decomposition.

¹⁸ This is probably the biggest advantage over conditional quantile methods like the one in Machado and Mata (2005), which is basically impossible to implement with reasonable computing resources for datasets larger than a few thousand observations.

functionals – e.g. quantiles – of the distribution. The influence function (IF) of the functional $\nu(G_W)$, for the underlying wage distribution $G_W(w)$ is defined as:

$$(11) \quad IF(w, \nu, G_W) \equiv \lim_{h \rightarrow 0} \frac{\nu(G_{W,h,\Delta_W}) - \nu(G_W)}{h} = \left. \frac{\partial \nu(G_{W,h,\Delta_W})}{\partial h} \right|_{h=0},$$

where $G_{W,h,\Delta_W} = (1-h)G_W + h\Delta_W$ is the perturbed wage distribution with the point mass distribution Δ_W at w . We are in general interested in what happens to the functional not in the case of a point-mass perturbation, but if the distribution G_W moves towards a new distribution, G_W^* . For that, we need the directional derivative of $\nu(G_W)$, in the direction of G_W^* , or, a function called the integrated influence function by Cowell and Victoria-Feser (1993).

$$(12) \quad \begin{aligned} IIF(\nu, G_W) &\equiv \left. \frac{\partial \nu(G_{W,h,G_W^*})}{\partial h} \right|_{h=0} = \int IF(w, \nu, G_W) d(G_W^* - G_W)(w) = \\ &= \int IF(w, \nu, G_W) dG_W^*(w), \end{aligned}$$

where the perturbed distribution is now $G_{W,h,G_W^*} = (1-h)G_W + hG_W^*$, and the last equality holds because $\int IF(w, \nu, G_W) dG_W(w) = 0$, by definition. The IIF's interpretation as a directional derivative is important intuitively, since it is suitable to determine the approximate value that functional $\nu(\cdot)$ takes on when $G_W(w)$ is perturbed by h times $G_W^*(w)$. More formally, using a local first-order Taylor approximation:

$$(13) \quad \nu(G_{W,h,G_W^*}) \approx \nu(G_W) + h \cdot IIF(w, \nu, G_W)$$

Now define the recentered influence function (RIF) by adding back the IF to the original functional of the distribution, that is,

$$(14) \quad RIF(w, \nu, G_w) \equiv \nu(G_w) + IF(w, \nu, G_w).$$

The RIF has the convenient feature that its expectation is equal to $\nu(G_w)$. Moreover, it is easy to show that (12) also holds for the RIF, such that

$$(15) \quad \frac{\partial \nu(G_w, h; G_w^*)}{\partial h} \Big|_{h=0} = \int RIF(w, \nu, G_w) d(G_w^* - G_w)(w).$$

Now assume that the perturbation is caused by a change in the distribution of some underlying worker characteristics. The unconditional distribution of wages can be expressed as in (1) in terms of the conditional distribution of wages given the covariates, and the marginal distribution of the covariates as

$$(16) \quad G_w(w) = \int G_{w|X}(w|X) dG_X(X).$$

Again, assuming invariance of the conditional distribution to changes in the distribution of X , the perturbed (counterfactual) distribution of wages is given by

$$(17) \quad G_w^*(w) = \int G_{w|X}(w|X) dG_X^*(X),$$

where the perturbation in $G_w^*(w)$ is now due to a perturbation in the distribution of the covariates, $G_X^*(X)$. By plugging into (15), and applying the law of iterated expectations we get

$$(18) \quad \frac{\partial v(G_{w,h,G_w^*})}{\partial h} \Big|_{h=0} = \int \mathbb{E}[RIF(w, v, G_w) | X = x] d(G_X^* - G_X)(x).$$

Remember that X stands for a vector of k covariates, $[X_t]_{t=1}^k$, distributed as G_X . Consider now a ceteris paribus location shift in the j th covariate, so that the new set of covariates X^h is equal to $[X_t^h]_{t=1}^k$, $X_t^h = X_t$ for $t \neq j$, and $X_t^h = X_t + h$ for $t = j$. Let the distribution of X^h be denoted by the perturbed distribution, G_X^* . It can be derived that the unconditional distribution of wages will then be $G_{w,h,G_w^*} = (1-h)G_w + hG_w^*$. The central theorem in Firpo et al. (2009) states that the local effect of a location shift in one of the covariates on some functional of the unconditional distribution of wages, keeping the other covariates constant, is given by

$$(19) \quad \frac{\partial v(G_{w,h,G_w^*})}{\partial h} \Big|_{h=0} = \int \frac{\partial \mathbb{E}[RIF(w, v, G_w) | X = x]}{\partial x_j} dG_X(x).$$

In other words, one can express the ceteris paribus effect of the change in the distribution of a covariate on any distributional statistic by the average partial effect of that covariate on the conditional expectation of its recentered influence function. So once a functional form is specified for the conditional expectation of the RIF, usual regression methods can be used to estimate the average partial effect. Firpo et al. (2009) discuss in detail the choice of the functional form; in this paper, I will assume that the conditional expectation is linear – that is $\mathbb{E}[RIF(w, v, G_w) | X = x] = X\beta$ –, but I will account for the possible specification error when performing the decomposition.

It is easy to see that to estimate the effect of changes in the distribution of worker characteristics on some functional of the unconditional distribution of wages – e.g. the variance, the median or other quantiles –, one only has to determine the RIF corresponding to

that particular functional at every observed wage in the sample, and regress the RIF values on individual characteristics. For example, it is possible to answer the question how the increase in the share of workers employed by foreign-owned companies affected certain quantiles of the unconditional wage distribution.

The RIF of the τ th quantile, q_τ , of the unconditional distribution of wages is calculated as

$$(20) \quad RIF(w, q_\tau, G_w) = q_\tau + \frac{\tau - \mathbb{I}\{w \leq q_\tau\}}{g_w(q_\tau)},$$

where $\mathbb{I}\{\cdot\}$ is an indicator function. An estimate for (20) is obtained by estimating q_τ by the sample quantile, and the corresponding density, $\hat{g}_w(q_\tau)$, by a kernel density estimator. Note that due to the presence of the indicator function in (20), in case of quantiles, assuming a linear functional form for $\mathbb{E}[RIF(w, q_\tau, G_w)|X]$ effectively means estimating a linear probability model, where the dependent variable is the estimated RIF and the covariates are individual characteristics of interest.¹⁹

With $\mathbb{E}[RIF(w, q_\tau, G_w)|X]$ specified, it is straightforward to adapt the standard OB decomposition framework to the RIF regressions. As noted earlier in the section, wages in year t are generated by the wage structure function $w_t = s_t(X, \varepsilon)$. The standard OB framework hinges on the assumption of linear additive separability, that is, $w_t = s_t(X, \varepsilon) = X_t\beta_t + \varepsilon_t$. Then the overall change in the unconditional mean of wages between years e and b , $\Delta_Q^t = \mathbb{E}(w|t = e) - \mathbb{E}(w|t = b)$, is equal to $\mathbb{E}(X|t = e)\beta_e - \mathbb{E}(X|t = b)\beta_b$ under the additional assumption of $\mathbb{E}(\varepsilon_t|X) = \mathbf{0}$, and by applying the law of iterated expectations. The difference can be then decomposed by adding

¹⁹ Firpo et al. (2009) discuss alternative estimation methods of the average partial effects, such as the logit and a non-parametric method. As noted earlier, I will assume that the LPM estimates the partial effects consistently. Note that it is necessary to assume linearity for carrying out the detailed decomposition, but I will account for possible specification errors by reweighting.

and subtracting the counterfactual expected wage when workers in year b earn the returns that prevail in year e :

$$(21) \quad \Delta_Q^t = [\mathbb{E}(X|t = e) - \mathbb{E}(X|t = b)]\beta_e + \mathbb{E}(X|t = b)(\beta_e - \beta_b) =$$

$$= \Delta_C^t + \Delta_S^t,$$

where the first term is the aggregate composition (or explained) effect and the second term is the aggregate wage structure (or unexplained) effect. Estimates for the composition and wage structure effects are obtained by replacing $\mathbb{E}(X|t)$ -s by sample averages, and β_t -s by OLS estimates.

Returning to the RIF quantile regression framework, the argument goes along the same lines, but the starting points are modified to the overall effect $\Delta_Q^{q\tau} = \mathbb{E}(RIF(w, q_\tau, G_w)|t = e) - \mathbb{E}(RIF(w, q_\tau, G_w)|t = b)$, and to the data generating process $RIF_t(w, q_\tau, G_w) = X_t\beta_t^{(q\tau)} + \varepsilon_t^{(q\tau)}$, for every quantile of rank τ in year t . The expectations and the coefficients can be estimated the same way as in the case of the OB method.

Barsky et al. (2002) pointed out that the classical OB decomposition will be inconsistent if the expectation $\mathbb{E}(w|X, t)$ is not linear. This is also true in case of the RIF decomposition. Fortin et al. (2011) propose to use the DFL reweighting method to account for possible specification biases in the decomposition. To see this, first let us construct the counterfactual distribution of wages when the distribution of year e characteristics is reweighted so that it resembles the distribution in year b . That is, we need

$$(22) \quad G_{w,e}^*(w) = \int G_{w|X,e}(w|X) \Psi^e(X) dG_{X,e}(X),$$

where the reweighting factor is shown by DiNardo et al. (1996) to be

$$(23) \quad \Psi^e(X) = \frac{dG_{X,b}(X)}{dG_{X,e}(X)} = \frac{Pr(X|t=b)}{Pr(X|t=e)} = \frac{Pr(t=b|X) \cdot Pr(t=b)}{Pr(t=e|X) \cdot Pr(t=e)}.$$

The last equality follows from Bayes' Law. The estimate, $\hat{\Psi}^e(X)$, is obtained by pooling data from the base and end periods, and estimating a binary outcome model for either $Pr(t = b|X)$, or for $Pr(t = e|X)$, and multiplying the predicted conditional probabilities by the share of observations in the respective period.²⁰ The counterfactual expectation of wages is then estimated by sample averages in the reweighted sample of year e , multiplied by OLS estimates of labor market returns from a reweighted regression of RIF values on covariates, that is, $\mathbb{E}^*(X|t = e)\beta_e^* \xrightarrow{est} \bar{X}_e^* \hat{\beta}_e^*$.

The decomposition is now given by the difference between actual averages and the reweighted counterfactual average as

$$(24) \quad \Delta_{\mathcal{O}}^{qr} = \Delta_{\mathcal{C}}^{qr} + \Delta_{\mathcal{S}}^{qr} = (\bar{X}_e \hat{\beta}_e^{(r)} - \bar{X}_e^* \hat{\beta}_e^{*(r)}) + (\bar{X}_e^* \hat{\beta}_e^{*(r)} - \bar{X}_b \hat{\beta}_b^{(r)}),$$

which can be further decomposed into “true” composition and wage structure effects and error terms by

²⁰ Obviously, since the subsamples for the two periods are mutually exclusive, $Pr(t = b|X) = 1 - Pr(t = e|X)$, and $Pr(t = b) = 1 - Pr(t = e)$.

$$(25) \quad \Delta_C^{qr} = (\bar{X}_e - \bar{X}_e^*)\hat{\beta}_e^{(r)} + (\hat{\beta}_e^{(r)} - \hat{\beta}_e^{*(r)})\bar{X}_e^*, \text{ and}$$

$$(26) \quad \Delta_S^{qr} = \bar{X}_b(\hat{\beta}_e^{*(r)} - \hat{\beta}_b^{(r)}) + (\bar{X}_b - \bar{X}_e^*)\hat{\beta}_e^{*(r)}.$$

In both expressions, the first term represents the pure composition/wage structure effect. Within (25), $(\hat{\beta}_e^{(r)} - \hat{\beta}_e^{*(r)})\bar{X}_e^*$ reflects the specification (or approximation) error that arises if the conditional expectation of the RIF is not linear, and it also captures errors from the fact that RIF regressions are based on local approximations of unconditional wage effects. If the specification error is found to be small, it is indicative of the RIF regression doing a good job in estimating the effects of large shifts in the distribution of covariates. It is especially important to check this term in the context of this paper, since the share of workers employed by foreign firms increased at a high rate over the years. $(\bar{X}_b - \bar{X}_e^*)\hat{\beta}_e^{*(r)}$ in Δ_S^{qr} is called the reweighting error as it shows how well the reweighted distribution of characteristics in year e mimics the distribution in year b . If $\hat{\Psi}^e(X)$ was estimated consistently, this term tends to be close to zero in large samples.

Since the focus of this paper is the effect of foreign ownership, I am interested in the contribution of the foreign ownership dummy in X to the total composition and wage structure effects. Because of the additive separability assumption, the contributions of single covariates can be partialled out easily just like in the case of the OB decomposition. The pure composition effect in (25) can be decomposed in detail into

$$(27) \quad (\bar{X}_e - \bar{X}_e^*)\hat{\beta}_e^{(r)} = (\bar{f}_e - \bar{f}_e^*)\hat{\beta}_{1,e}^{(r)} + \sum_{l=2}^k (\bar{Z}_{l,e} - \bar{Z}_{l,e}^*)\hat{\beta}_{l,e}^{(r)},$$

and the pure wage structure effect in (26) into

$$(28) \quad \bar{X}_b(\hat{\beta}_s^{(\tau)} - \hat{\beta}_b^{(\tau)}) = (\hat{\beta}_{0,s}^{(\tau)} - \hat{\beta}_{0,b}^{(\tau)}) + \bar{f}_b(\hat{\beta}_{1,s}^{(\tau)} - \hat{\beta}_{1,b}^{(\tau)}) + \\ + \sum_{l=2}^k \bar{Z}_{l,b}(\hat{\beta}_{l,s}^{(\tau)} - \hat{\beta}_{l,b}^{(\tau)}).$$

$(\hat{\beta}_{0,s}^{(\tau)} - \hat{\beta}_{0,b}^{(\tau)})$ represents the wage structure effect for the omitted group, the terms involving elements of \bar{Z} capture the contributions of individual characteristics other than ownership, while the two contributions of main interest are $(\bar{f}_s - \bar{f}_s^*)\hat{\beta}_{1,s}^{(\tau)}$ and $\bar{f}_b(\hat{\beta}_{1,s}^{(\tau)} - \hat{\beta}_{1,b}^{(\tau)})$. The former measures how the change in foreign penetration contributed to the overall change in the value of a particular quantile of the unconditional wage distribution between the end and the base period, while the latter estimates the contribution of changes in the returns to being employed at a foreign-owned company.

5. Stylized Facts on Changes in the Wage Distribution

Before turning to the decomposition analysis of FDI and wage inequality, I present some descriptive figures about the evolution of the dispersion of earnings in general. Figure 2 follows changes in real wages over time at five selected percentiles of the wage distribution, compared to their 1986 value. Median earnings have declined dramatically right at the beginning of transition, and even though they started to recover in 1992, the stabilization package introduced in 1995 caused wages to fall back to the 1992 level. Lower quantiles were hit more heavily by the shocks of transition and stabilization, with workers at the lowest decile earning in 1997 just 70 percent of what workers at the same decile had earned in 1986. The tide turned in 1997, when real earnings at all points of the distribution started to rise, and have been on the rise basically until the last sample year. However, since the relapse in the early nineties was so radical, 2002 was the first year when median wages reached again their pre-transition level.

Figure 2 also provides evidence on the diverging patterns of real wage changes at different points of the distribution – that is, on increasing wage inequality –, a tendency that started in 1992 and has been maintained more or less throughout the whole sample period. The distance between the top and the bottom decile is the largest in 2000, but this seems to be partly a consequence of the strange behavior of wages at the tenth percentile. The surprising increase in the latter from 2000 to 2002 is due to a government intervention that left the real value of the minimum wage in 2002 seventy percent of its value in 2000, and the minimum wage became the tenth percentile of the distribution. The decomposition analysis that follows later refers only to the 1992-2000 period, so those results will not be affected by changes in the minimum wage legislation.

The evolution of wage inequality is easier to follow on Figure 3, where dispersion is measured by the variance of log wages. With a short break in 1995, inequality grew substantially and at a fast pace from the end of the socialist regime to 2000, reaching more than 180 percent of its pre-transition level. The huge increase in the minimum wage between 2000 and 2002 also affected the variance, but the upward sloping pattern continued after 2002 for four years. Finally, in the last two years of data, the dispersion decreased, especially in 2008. Figure 3 also reflects differences in the nature of wage dispersion between the beginning and the end of the period by applying a standard variance decomposition to divide total variance into a within-firm and a between-firm component. While during the last years of socialism, inequality was almost completely a within-firm phenomenon, with the introduction of market mechanisms in wage determination, the share of between-firm variation increased to explain close to sixty percent of total variation in wages by the end of the period. Moreover, between-firm variation seems to closely follow the evolution of the overall variance: it is rising when total dispersion is rising, and falling when the spread in the

whole distribution is also falling. This suggests that changes in inequality are mostly linked to factors that are correlated with firm-level heterogeneity.

In all what follows, I will focus on the period between 1992 and 2000, and I will consider the market for female and male employees as separate labor markets. In Figure 4, I show how the difference between the ninth and the first decile of the log wage distribution evolved for women and men. As represented by the solid line, the log 90-10 differential grew by 30 percent for both genders. The dashed curve right below shows that there is a difference in the location of the wage distributions of workers employed by foreign and domestic firms, since when removing average wages by ownership group, inequality decreases. This difference in the averages is growing over time, more strongly for men than for women. The evolution of sample averages of earnings is represented separately in Figure 5 as well. The two bottom curves in Figure 4 show that although the foreign wage premium is correlated with individual characteristics, region and industry, it is not completely explained by them. Residual wage inequality within groups of workers defined by education, experience, occupation, region and industry is much lower than total (unconditional) inequality but these factors do not account completely for either the level or for the growth in wage dispersion. When including the foreign dummy into the yearly wage regressions to predict residual wages, residual inequality further decreases, but residual earnings at the 90th percentile are still 2.9 times higher than at the 10th percentile for women (corresponding to a difference of 1.05 log points), and 3.3 times higher for men. The growth in residual inequality from 1992 to 2000 in the all inclusive setup is 21.5 and 22.9 percent, respectively.

Figure 6 demonstrates that with the exception of the early years in case of women, not only the mean (see Figure 5), but the variance is also larger for foreign-owned employees.²¹ This gap, however, does not exhibit a strongly increasing trend similar to the case of the

²¹ The coefficient of variation, defined as the standard deviation of log wages divided by the mean of log wages, is also higher for workers of foreign firms by 2-5%.

mean for the male distribution, but widens until 1999 for women. As we see, both the mean and the variance are higher for the foreign group, and both of these facts mean that an increase in FDI will result in a wider spread in the overall wage distribution. This may of course not necessarily be a direct effect of FDI itself, but also of some other factors correlated with ownership, but I will partial out the foreign effect in the next section.

Combining the differences in means and variances, it is useful to quantify how the between-group and within-group differences in wages between foreign and domestic workers and the relative share of each group in total employment contribute to the total variation in wages. Table 2a decomposes levels of variances at the beginning and at the end of the period, while Table 2b decomposes changes in the variance over time with the help of a rough-and-simple within-/between-group variance decomposition, where the groups are defined according to ownership status. The share of between-group variation increased heavily for both genders, partly because the group-level average wages diverged (i.e. the between-group variance increased), and partly because the share of foreign firms grew substantially in total employment. Of course, the main component in both the level and the change of the variance is the within term, which is not surprising when sorting the workforce into just two broad and heterogeneous groups, and without conditioning on any other variable. Again, for the within variance, the increasing share of the higher-variance foreign group is a factor, but not as much as in the case of the between-group term. Considering gender differences, all types of variation grew more strongly for men, and composition effects have a higher importance for them than for women. For men, the share of foreign employment in 2000 was 6.8 times higher than in 1992, while the same ratio is 6.2 for women, although for the latter group, employees of foreign-owned companies are more numerous in relative terms in both years.

In the next section, I will use more sophisticated decomposition techniques to estimate the ceteris paribus contribution of FDI to changes in the shape of the unconditional wage distribution.

6. Decomposition Results

Before presenting the results of a parametric decomposition by quantile, Figures 7, 8a and 8b help to visually inspect how changes in FDI are correlated with changes in the shape of the density of wages, using the non-parametric density decomposition described in Section 4.1. In Figure 7, I depicted kernel density estimates of actual wage distributions in 1992 and 2000, and for the counterfactual distribution of a hypothetical state where all characteristics of workers are distributed according to actual shares in 2000, except for ownership that is supposed to be distributed as in 1992; and all returns to individual characteristics are determined by the 2000 wage structure (including returns to foreign ownership). The construction of this counterfactual density followed equation (4) by reweighting the 2000 sample with weights defined in (5). The difference between the counterfactual distribution (marked with the dashed line) and the actual distribution in 2000 shows the effect of the change in FDI penetration between 1992 and 2000, given that the identifying assumptions outlined in Section 4.1 hold. Since the 2000 form of the conditional distribution of wages given the covariates is supposed to prevail, this difference should only capture the effect of the change in the distribution of foreign ownership (i.e. in the share of foreign employment), as opposed to changes in the foreign wage premium.

For both men and women, the changes in factors other than ownership composition increase inequality in the lower end of the distribution, while they do not seem to alter much the shape of the density in the upper end. In contrast, the increase in FDI shifts the location of most of the quantiles to the right, with the exception of very low wages. Besides the

location shift, we can observe an increase in the spread of the distribution, especially in case of male earnings.

Figures 8a and 8b display the wage structure effects of foreign ownership. As discussed in Section 4.1, isolating this effect on the density is trickier, and one has to construct the necessary counterfactual distributions separately for the base and end periods, and the effect of the change in the foreign wage premium is given by a difference-in-differences formula in (10). The counterfactual distribution in this case is one where the sample of workers of domestic companies is reweighted in each year to simulate a state where no workers are employed by foreign-owned firms, and the distribution of characteristics in the whole sample is represented by the reweighted subsample of domestic firms. We see in Figure 8a that foreign ownership basically has no effect in the base year through a different pay scheme on the density of wages. This is not surprising of course, given that the fraction of foreign-owned enterprises was anyway very small in 1992. In 2000, the picture changes slightly – more so for women than for men – since higher wages paid by foreign employers shift the unconditional earnings distribution to the right. For men, this seems to be a pure location shift – with the exception of the bottommost percentiles – but for women, it also results in a higher kurtosis of the density. Taken all density graphs together, we can conclude that the composition effect of FDI on the shape of the wage distribution was the dominant factor and not the wage structure effect. In what follows, I will estimate these effects parametrically for selected quantiles of the unconditional wage distribution.

The rest of this section builds on the methodological framework introduced in Section 4.2. The steps of estimation are the following. First, I reweighted the sample in year 2000 by an estimated version of the weighting factor defined in (23). The purpose of this was to create the counterfactual distribution for the decomposition in which the distribution of characteristics mimics the distribution in 1992. Second, I estimate for 19 quantiles (for every

fifth percentile from the 5th to the 95th) the RIF as defined by (20). Third, I use the estimated RIFs as dependent variables in the unconditional quantile regressions that I run for every quantile separately, and where the regressors are foreign ownership and other covariates including education, potential experience, occupation, region and industry. The regressions are estimated by OLS clustering for firm-level heteroskedasticity. Finally, I decompose changes in RIFs (basically changes in unconditional quantiles) with the standard Oaxaca-Blinder technique, but within the framework outlined in (24)-(28).

The estimated coefficients from the unconditional quantile regressions are presented in Figure 9a through Figure 10c, separately by gender (Figure 9 – men, Figure 10 – women) for the base year (gray lines and markers) and for the end year (black lines and markers). The coefficients are plotted against quantiles for each group of covariates. Remember from Section 4.2 that every coefficient measures the estimated effect of increasing the share of workers in the population with the given characteristic on the given quantile of the unconditional wage distribution, holding everything else constant. For example, the plot for university (and college) graduates in Figure 9a tells us that in 2000, a ten-percentage-point increase in the number of male workers with completed higher education would shift the 90th percentile of the unconditional wage distribution of men to the right by 16.5 percent ($0.1 * 1.615$), while the median by only 4 percent ($0.1 * 0.404$), all other things equal. It follows that positively sloped curves are evidence of a potentially inequality enhancing covariate, while characteristics with decreasing curves might attenuate wage inequality, if their relative importance increases in the population.

Although it would be instructive to spend time with the in-depth discussion of other covariates, since the focus of this paper is foreign ownership, I will concentrate on the foreign

effect.²² Consider first the case of men in Figure 9a. In 1992, the effect of a location shift in the share of foreign workers is clearly estimated to be dispersion enhancing. Nonetheless, the foreign effect is positive at all quantiles, so that higher FDI inflow has a wage increasing impact not only on average, but at every point of the distribution. Note, however, that this does not mean that all workers – including those of domestic firms – necessarily benefit since one of the crucial assumptions of the RIF regression framework is the invariance of the conditional wage distribution to underlying changes in covariates, which effectively means that spillover effects of FDI are assumed to be zero. By 2000, the effects below the median get roughly equalized with the exception of the lowest quantile, but that coefficient is probably estimated with noise as apparent from the other plots as well. Above the median, the inequality enhancing effect remains about the same as it was in 1992.

For women, the pattern is more interesting, displaying a double U-shape. The relationship between the unconditional wage effect and the position in the distribution is concave in the bottom part of the distribution, while it is convex in the top part, with the inflection point being somewhere around the sixth decile in 1992 and around the median in 2000. In other words, a *ceteris paribus* growth in the share of female workers employed by foreign firms has an inequality increasing impact on very low and very high earnings, while an inequality mitigating impact on medium earnings. Also, the wage effects are estimated to be higher in 2000 than in 1992 for the low and the high end, but lower for the middle section of the distribution. Moreover, the largest effects are close to the median worker in 1992, and even in 2000, the effects close to the second to third deciles are as high as for the top two quantiles, which tells a completely different story from the case of men.

The estimated effects discussed so far only reflect a snapshot in time, but do not help *per se* in quantifying how actual changes in the distribution of covariates affected actual

²² Appendix Table 2 collects estimated foreign effects and significance levels from the unconditional quantile regressions, for nine deciles of the female and male wage distributions in 1992 and 2000.

changes in the unconditional distribution of wages over time. It is useful to refer to a remark by Fortin et al. (2011) who point out the importance of decomposition exercises by drawing attention to the fact that although numerous studies find the returns to education to be large (and significant), the differences in human capital endowment over time or across countries only account for a small part of either the growth in GDP over time or of GDP differences between countries. Thus, estimating large positive wage effects of foreign ownership at any point in time does not necessarily mean that a more pronounced presence of foreign investors in the Hungarian business sector had a large influence on the wage distribution. How large it was can be answered by the detailed decomposition of quantile influence functions.

Prior to the analysis of FDI's contribution to wage changes, I summarize the results of the aggregate decomposition according to (24)-(26) in Figure 11. Actual changes in wages by quantile follow an increasing function, indicating a rise in inequality across the distribution. The curves are very similar for the two genders, with the male plot being somewhat steeper in the middle. The two error terms – $(\hat{\beta}_e^{(t)} - \hat{\beta}_e^{*(t)})\bar{X}_e^*$ and $(\bar{X}_b - \bar{X}_e^*)\hat{\beta}_e^{*(t)}$ in (25) and (26), respectively – both lie very close to the zero line. The first one – representing specification or approximation error – being small means that, first, the bias arising from the linear specification of the conditional expectation of the RIF is small, and second, that although the RIF projection is a local approximation it is doing a good job in measuring the effect of larger changes in underlying characteristics.²³ The negligible magnitude of the second error term is evidence of the reweighted (counterfactual) distribution of covariates in 2000 being in fact close to the actual distribution of 1992, which is what we wanted.

By inspecting curves for the composition and wage structure components in Figure 11, we see that the positive real wage changes across the distribution are mostly driven by

²³ Note that we might have expected the specification error to be small due to the fully flexible specification of the RIF regressions by dummifying out educational, regional, industrial and experience categories.

composition effects, while wage structure effects are mostly responsible for the wider spread of the distribution by the end of the period, for both genders. Changes in the composition of the workforce do also have some inequality boosting impact, especially at the two ends of the distribution, but it is rather a location shift type of effect that dominates. Changes in the returns to skill and other characteristics affected negatively quantiles below the median, while benefitted the upper half. For the lowest ten percent of the distribution it was enough to offset the positive effects of composition changes so that wages at these quantiles decreased from 1992 to 2000.

Let us now move to the detailed decomposition by covariates as described in equations (27) and (28) for each quantile with rank τ . Figure 12 depicts the total contributions of each covariate (or group of covariates) to the decomposition, by adding up composition and wage structure effects for each. The “Other” plot merges the error terms from Figure 11 and the wage structure effect for the omitted group in the regressions, that is, the term $(\hat{\beta}_{0,s}^{(\tau)} - \hat{\beta}_{0,b}^{(\tau)})$ from equation (28). We see that foreign ownership is a major factor in wage changes over time for every quantile and for both genders. On average across quantiles, 0.085 log points are explained by the increase in FDI of the changes in male wages. This is quite substantial considering that wages changed by 0.143 log points on average. For women, the same numbers are 0.090 against a total change of 0.153. Turning to the by-quantile effects, the contribution of FDI to wage changes is mostly uniform across the male distribution, especially between the second and sixth deciles. Foreign ownership only has an inequality enhancing effect in the top forty percent of the distribution. Concerning female earnings, the pattern reminds us of the pattern of estimated coefficients from yearly regressions in Figure 10a. We observe the double U-shaped curve, only in a much flatter version. Regarding other characteristics, education and industry had strong inequality increasing effects between 1992 and 2000 for both genders, while regional changes decreased

dispersion. Occupational changes have a large and inequality mitigating effect for men, and are less important, and have a mixed effect on inequality for women.

Finally, I divide the total contributions of characteristics to effects due to changes in the distribution of the characteristic and due to changes in the returns to the characteristic. I will only consider the “pure” composition and wage structure effects and not deal with the specification and reweighting errors as these were shown to be plausibly negligible. Figures 13a and 13b clearly demonstrate that the dominant factor in the contribution of foreign ownership to wage changes is the increase in the share of employees working for foreign firms as wage structure effects are very close to zero for both men and women. Within composition effects, FDI is the leading explanation across the whole wage distribution. Since wage effects are negligible, the heterogeneity of the foreign composition effect by quantiles takes on the same shape as the total foreign effect in Figure 12. For men, the main impact of the rise in the prevalence of foreign companies was a location shift of the distribution, with extra wage premia above the median, increasing in the rank in the distribution. However, because the curve of total wage changes for the upper quantiles is even steeper, the relative explanatory power of foreign ownership decreases when moving towards higher earnings, as other factors’ importance – especially that of education – gets higher.

As Appendix Table 3 summarizes, the total change in the log 90-10 wage differential was 0.376 between 1992 and 2000 for men, and the implied contribution of foreign composition effects is only 0.020 log points (computed as the difference between the composition effects estimated on the ninth and the first deciles), that is, about five percent of the total change. For the 90-50 differential, it is 0.187 versus 0.034, while for the 50-10 it is 0.189 versus -0.013 log points. This means that for the lower half of the distribution, FDI even had an attenuating effect on growing inequality. For comparison, note that changes in the skill composition of the workforce, as measured by highest degree of education,

contributed by 0.040 log points to the total change in the 90-10 differential, while by 0.028 log points to the change in the 90-50 differential. Note also that composition changes related to labor market experience, regions and occupations had only a minuscule impact on changes in unconditional wages.

For women, the major factors regarding composition effects in the upper graph of Figure 13b are foreign ownership share, education, industry and occupation, with the first being by far the most important. The foreign effect varies more by quantile than in case of men, displaying an inequality decreasing pattern in the major middle part of the wage distribution, increasing dispersion only at the low and high ends. More precisely, the lower panel of Appendix Table 3 indicates that of the total changes in the 90-10, 90-50 and 50-10 inequality measures of 0.349, 0.170 and 0.180 log points, changes in foreign penetration account for a mere 0.010, 0.018 and -0.008 log points, respectively.

The estimated wage structure effects are harder to interpret as they depend on the choice of the omitted group. Firpo et al. (2011) discuss the issue in detail and survey possible solutions from the literature. The omitted group problem, however, does not affect binary variables, so the estimated coefficients on the foreign dummy would not change, if I specified another base group. As we saw, the contribution of FDI to overall wage changes by quantile through changes in the foreign wage premia is rather small. Wage structure effects of all other covariates have to be taken with care.

7. Conclusions

The novelty of this research was to examine the relationship between foreign direct investment and changes in the unconditional distribution of wages using a new method that enables the detailed decomposition of wage changes by quantiles on a sample of workers in the Hungarian business sector. Hungary experienced both a huge amount of FDI inflow during the nineties and significant changes in the location and the shape of the earnings

distribution, so it provides an excellent subject of analysis regarding the effects of foreign ownership. A further advantage was insured by linked employer-employee data that facilitated the exact measurement of ownership status and the inclusion of several worker characteristics to the analysis as controls.

I found that wage inequality has been on a rising path throughout the 1989-2008 period, with some breaks that however do not affect the subject period of the main focus of investigation, that is the years between 1992 and 2000 when the largest increase in wage dispersion and the largest growth in the share of foreign employment happened. For this latter interval of time, inequality rose by thirty percent for men and for women, as measured by the log 90-10 wage differential. At the same time, the share of workers employed by foreign-owned firms increased from five to close to forty percent.

A non-parametric decomposition of the change in the density of unconditional wages between 1992 and 2000 shows that it was mainly the change in the composition of workforce by ownership status and not a change in returns to ownership status through which FDI affected the wage distribution. Also, this effect was rather a shift in the location of the distribution and less of a change in the shape of the density function. Going beyond the visual inspection of densities, I applied the parametric method developed by Firpo et al. (2007, 2009) that builds on the recentered influence function to (i) estimate the effect of the prevalence of worker characteristics on distributional statistics, and to (ii) decompose changes at various points of the unconditional wage distribution into composition and wage structure effects of these characteristics, with particular attention to foreign employment status. The effects of an increase in FDI at any point in time are estimated to be large, positive and significant across the distribution for both genders. For men, the pattern of the effects is suggestive of an inequality enhancing impact, while for women, it is amplifying dispersion at the two ends of the distribution, but has an alleviating impact in the middle.

Concerning the contributions of FDI to actual wage changes by quantiles, the rise in the share of employees working for foreign firms had a substantial positive composition effect at every quantile, but the role of changes in foreign wage premia over time is negligible. The dominant effect of the composition change related to ownership is a general increase in wages across the whole distribution, but it also accounts for about twenty percent of the rise in male inequality above the median, while it explains ten percent for women. The growth in FDI had a slight inequality mitigating effect in the lower half of the distribution in case of both genders.

It is important to note that these estimated effects can only be considered as causal effects of FDI if the very stringent identifying assumptions set in the paper are met, which is probably not true. In particular, the method assumes away (i) different mechanisms of participation in the labor market in 1992 and 2000, (ii) self-selection of workers into foreign employment based on unobservable heterogeneity, (iii) foreign investors systematically selecting target firms for acquisitions with characteristics that are correlated with changes in the wage distribution, and (iv) wage spillover effects of FDI on workers' wages employed by domestic firms. These are all relevant concerns that are not addressed directly in the paper. However, I pointed out the importance of FDI and/or of factors correlated with FDI in examining changes in the distribution of unconditional earnings so that the paper moved beyond the typical analysis of the foreign effect on conditional average wages. Integrating solutions to the above listed identification issues is a subject of future research.

References

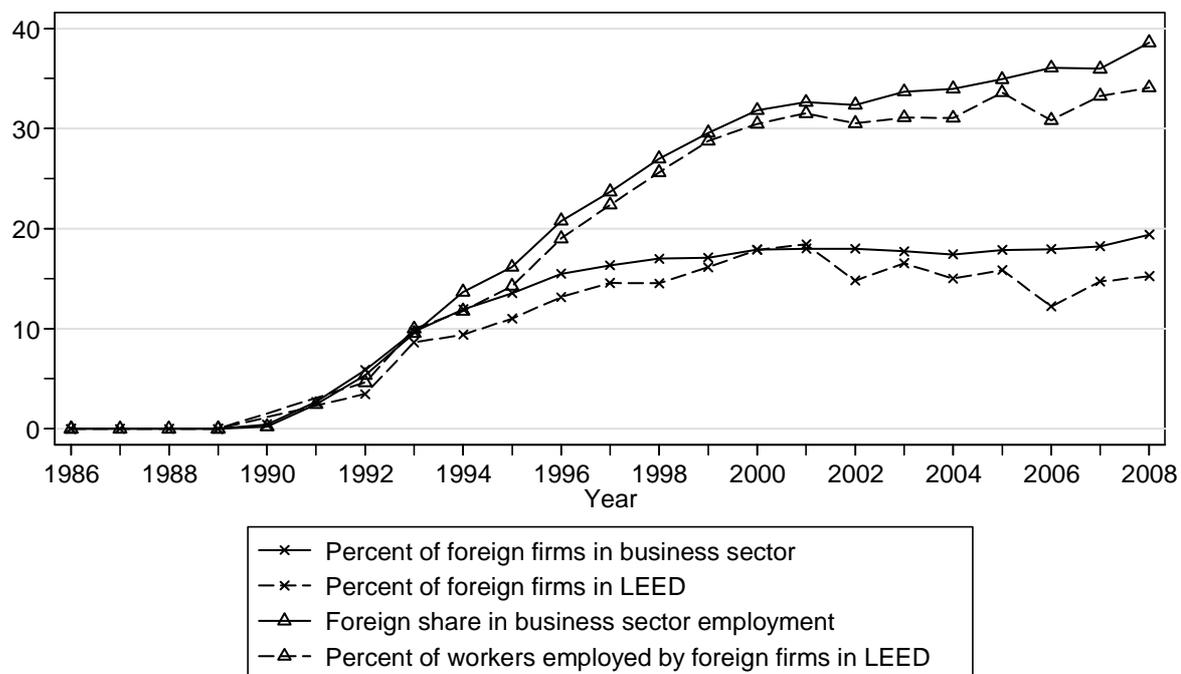
- Abowd, JM, Kramarz F, and Margolis DN. High Wage Workers and High Wage Firms. *Econometrica* 1999; 67; 251-334.
- Aitken B, Harrison A, and Lipsey RE. Wages and Foreign Ownership: A Comparative Study of Mexico, Venezuela and the United States. *Journal of International Economics* 1996; 40; 345-371.
- Almeida R. The Labor Market Effects of Foreign Owned Firms. *Journal of International Economics* 2007; 72; 75-96.
- Andrews M, Bellmann L, Schank T, and Upward R. The Takeover and Selection Effects of Foreign Ownership in Germany: An Analysis Using Linked Worker-Firm Data. Research Paper 2007/2008.
- Antal G. Dispersion of Earned Income in Transition: The Robustness of Earnings Inequality Growth in the Hungarian Business Sector, 1986-2005. Unpublished PhD dissertation chapter 2011, Department of Economics, Central European University.
- Autor DH, Katz LF, and Kearney MS. Trends in U.S. Wage Inequality: Revisiting the Revisionists. *Review of Economics and Statistics* 2008; 90; 300-323.
- Autor DH, Katz LF, and Kearney MS. Rising Wage Inequality: The Role of Composition and Prices. NBER Working Paper 2005; 11628.
- Blinder A. Wage Discrimination: Reduced Form and Structural Estimates. *Journal of Human Resources* 1973; 8; 436-455.
- Bound J, Johnson G. Changes in the Structure of Wages in the 1980s: An Evaluation of Alternative Explanations. *American Economic Review* 1992; 82; 371-392.
- Brown DJ, Earle JS, and Telegdy Á. The Productivity Effects of Privatization: Longitudinal Estimates from Hungary, Romania, Russia, and Ukraine. *Journal of Political Economy* 2006; 114; 61-99.
- Brown DJ, Earle JS, and Telegdy Á. Employment and Wage Effects of Privatization. *Economic Journal* 2010; 120; 683-708.
- Brown DJ, Earle JS. Understanding the Contributions of Reallocation to Productivity Growth: Lessons From a Comparative Firm-Level Analysis. Upjohn Institute for Employment Research Working Paper 2008; WP08-141.
- Canyon MJ, Girma S, Thompson S, and Wright PW. The Productivity and Wage Effects of Foreign Acquisitions in the United Kingdom. *The Journal of Industrial Economics* 2002; 50; 85-102.
- Cowell F, Victoria-Feser M-P. Robustness Properties of Inequality Measures. LSE STICERD Discussion Paper 1993; No. DARP 01.
- Davis SJ, Haltiwanger JC. Earnings Dispersion Between and Within U.S. Manufacturing Plants, 1963-86. *Brookings Papers on Economic Activity* 1991; 115-180.
- DiNardo J, Fortin NM, and Lemieux T. Labor Market Institutions and the Distribution of Wages, 1973-1992: A Semiparametric Approach. *Econometrica* 1996; 64; 1001-1044.
- DiNardo J, Lemieux T. Diverging Male Inequality in the United States and Canada, 1981-1988: Do Institutions Explain the Difference? *Industrial and Labor Relations Review* 1997; 50; 629-651.

- Dunne T, Foster L, Haltiwanger JC, and Troske KR. Wage and Productivity Dispersion in United States Manufacturing: The Role of Computer Investment. *Journal of Labor Economics* 2004; 22; 397-429.
- Earle JS, Telegdy Á, and Antal G. Inward Foreign Direct Investment and the Hungarian Wage Structure. Unpublished PhD dissertation chapter 2011, Department of Economics, Central European University.
- Eriksson T, Pytliková M, and Warzynski F. Increased Sorting and Wage Inequality in the Czech Republic: New Evidence Using Linked Employer-Employee Dataset. Aarhus University Working Paper 2009; 09-5.
- Fazekas K, and Koltay J (Eds). *The Hungarian Labour Market: Review and Analysis*. Institute of Economics, Hungarian Academy of Sciences: Budapest; 2002.
- Feliciano Z, Lipsey RE. Foreign Ownership, Wages, and Wage Changes in U.S. Industries, 1987–92. *Contemporary Economic Policy* 2006; 24; 74–91.
- Firpo S, Fortin N, and Lemieux T. Decomposing Wage Distributions using Recentered Influence Function Regressions. Unpublished manuscript 2007; University of British Columbia. Available at <http://faculty.arts.ubc.ca/nfortin/research.html>; last accessed 30 June 2011.
- Fortin N, Lemieux T, and Firpo S 2011. Decomposition Methods in Economics. In: Ashenfelter O, Card D (Eds), *Handbook of Labor Economics*, Volume 4A, Elsevier: Amsterdam; 2011. p. 1-102.
- Frydman R, Rapaczynski A, and Earle JS (Eds). *The Privatization Process in Central Europe*, Central European University Press: Budapest; 1993.
- Ganguli I, Terrell K. Institutions, Markets and Men's and Women's Wage Inequality: Evidence from Ukraine. *Journal of Comparative Economics* 2006; 34; 200-227.
- Girma S, Görg H (2007). Evaluating the Foreign Ownership Wage Premium Using a Difference-in-Differences Matching Approach. *Journal of International Economics* 2007; 72; 97-112.
- Görg H, Strobl E, and Walsh F. Why Do Foreign-Owned Firms Pay More? The Role of On-the-Job Training. *Review of World Economics* 2007; 143; 464-482.
- Haltiwanger JC, Lane JJ, and Spletzer JR. Wages, Productivity, and the Dynamic Interaction of Businesses and Workers. *Labour Economics* 2007; 14; 575-602.
- Hampel FR. The Influence Curve and Its Role in Robust Estimation. *Journal of the American Statistical Association* 1974; 69; 383-393.
- Heyman F, Sjöholm F, and Gustavsson Tingvall P. Acquisitions, Multinationals and Wage Dispersion. IFN Working Paper 2006; No. 675.
- Heyman F, Sjöholm F, and Gustavsson Tingvall P. Is there Really a Foreign Ownership Wage Premium? Evidence from Matched Employer-Employee Data. *Journal of International Economics* 2007; 73; 355-376.
- Huttunen, K. The Effect of Foreign Acquisition of Employment and Wages: Evidence from Finnish Establishments. *The Review of Economics and Statistics* 2007; 89; 497-509.
- Jolliffe D, Campos N. Does Market Liberalization Reduce Gender Discrimination? Econometric Evidence from Hungary, 1986-1998. *Labour Economics* 2005; 12; 1-22.
- Juhn C, Murphy KM, and Pierce B. Wage Inequality and the Rise in Returns to Skill. *Journal of Political Economy* 1993; 101; 410-442.

- Katz LF, Autor DH 1999. Changes in the Wage Structure and Earnings Inequality. In: Ashenfelter O, Card D (Eds), *Handbook of Labor Economics*, Elsevier: Amsterdam; 1999. p. 1463-1555.
- Katz LF, Murphy KM. Changes in Relative Wages, 1963-1987. *Quarterly Journal of Economics* 1992; 107; 35-78.
- Kaufmann D, Kraay A, and Mastruzzi M. Governance Matters III: Governance Indicators for 1996-2002. World Bank Policy Research Working Paper 2003; No. 3106.
- Keane MP, Prasad ES. Changes in the Structure of Earnings During the Polish Transition. *Journal of Development Economics* 2006; 80; 389-427.
- Kertesi G, Köllő J. Real Wages and Earnings Inequality, 1986-1996. (in Hungarian: Reálbérek és kereseti egyenlőtlenségek, 1986-1996) *Közgazdasági Szemle* 1997; 44; 612-634.
- Kertesi G, Köllő J. Economic Transformation and the Revaluation of Human Capital: Hungary, 1986-1999. *Research in Labor Economics* 2002; 21; 235-273.
- Kertesi G, Köllő J. Industrial Wage Differentials in Hungary, Pt. 1. (in Hungarian: Ágazati bérkülönbségek Magyarországon, 1. rész: Az ágazati járadékképződés alternatív modelljei.) *Közgazdasági Szemle* 2003a; 50; 923-938.
- Kertesi G, Köllő J. Industrial Wage Differentials in Hungary, Pt. 2. (in Hungarian: Ágazati bérkülönbségek Magyarországon, 2. rész: Járadékokon való osztozkodás koncentrált ágazatokban, szakszervezeti aktivitás jelenlétében.) *Közgazdasági Szemle* 2003a; 50; 1049-1074.
- Köllő J, Nagy Gy. Earnings Gains and Losses from Insured Unemployment in Hungary. *Labour Economics* 1996; 3; 279-298.
- Kornai J. *The Socialist System*. Oxford University Press: Oxford; 1992
- Lemieux T. Increasing Residual Wage Inequality: Composition Effects, Noisy Data, or Rising demand for Skill? *The American Economic Review* 2006; 96; 461-498.
- Lemieux T. The Changing Nature of Wage Inequality. *Journal of Population Economics* 2008; 21; 21-48.
- Lipsey RE, Sjöholm F. Foreign Direct Investment, Education and Wages in Indonesian Manufacturing. *Journal of Development Economics* 2004; 73; 415-422.
- Machado JF, Mata J. Counterfactual Decomposition of Changes in Wage Distributions Using Quantile Regressions. *Journal of Applied Econometrics* 2005; 20; 445-465.
- Martins P. Do Foreign Firms Really Pay Higher Wages? Evidence from Different Estimators. IZA Discussion Paper 2004; No. 1388.
- Martins P, Esteves LA. Foreign Ownership, Employment and Wages in Brazil: Evidence from Acquisitions, Divestments and Job Movers. IZA Discussion Paper 2008; No. 3542.
- Mihályi P. On the Quantitative Aspects of Hungarian Privatization. *Comparative Economic Studies* 1997; 39; 72-95.
- Neumann L. Does Decentralized Collective Bargaining Have an Impact on the Labor Market in Hungary? *European Journal of Industrial Relations* 2002; 8; 11-31.
- Oaxaca R. Male-Female Wage Differentials in Urban Labor Markets. *International Economic Review* 1973; 14; 693-709.
- OECD Reviews of Foreign Direct Investment: Hungary. OECD: Paris; 2000.

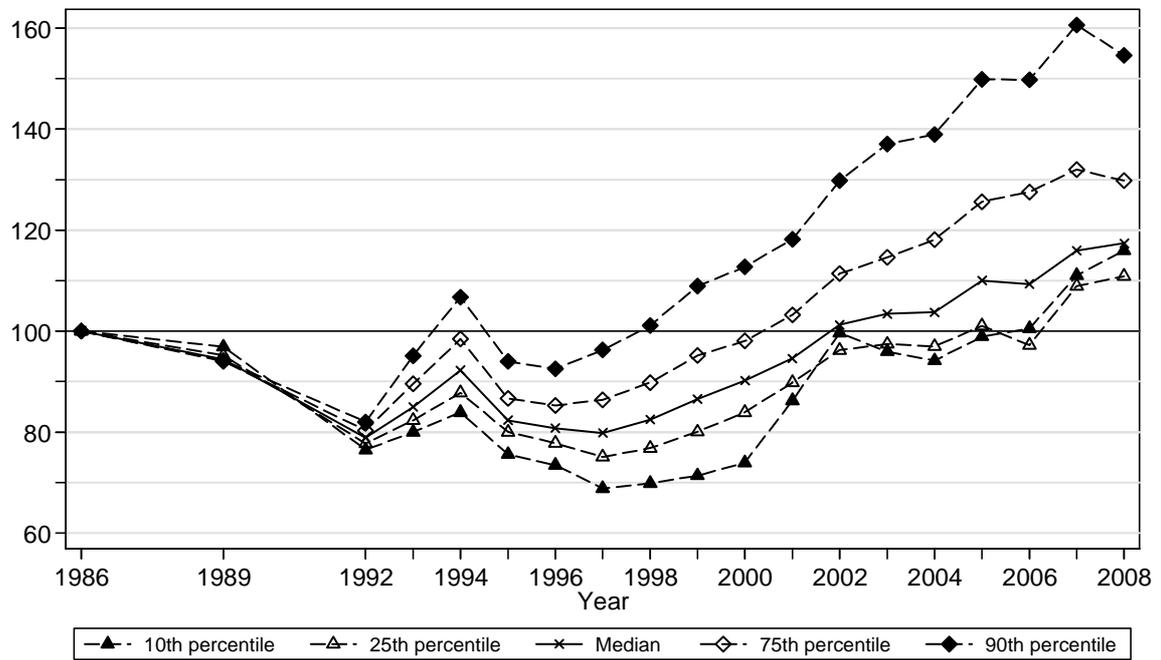
Tables and Figures

Figure 1: Foreign Penetration in the LEED and in the Business Sector, 1986-2008



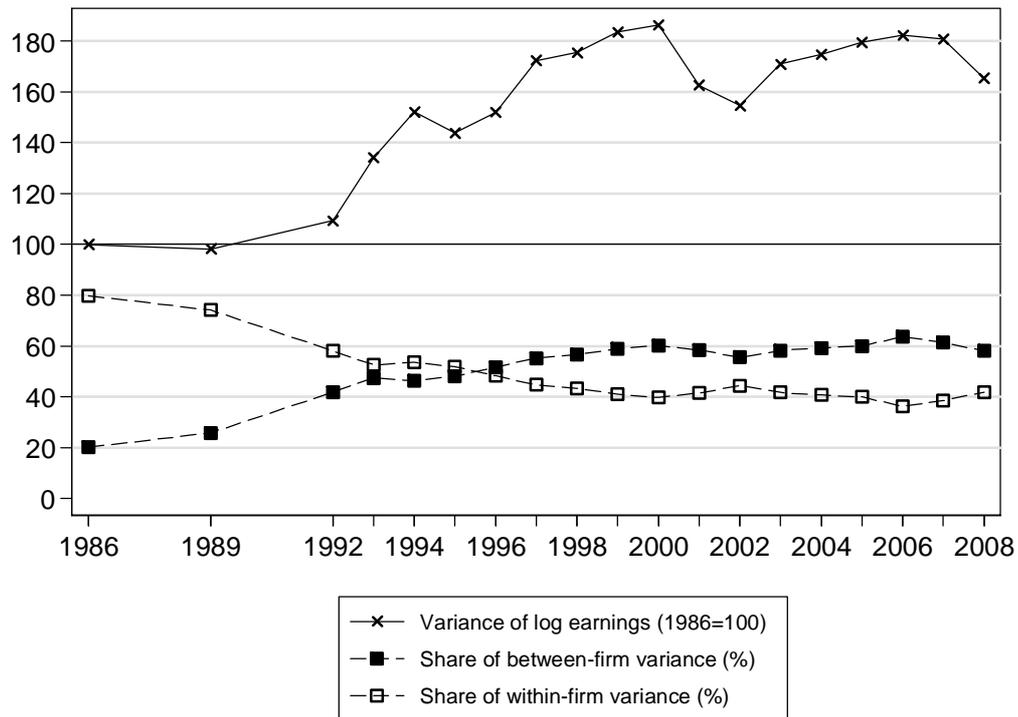
Notes: Only firms with more than 20 employees. Business sector shares are computed from a comprehensive administrative dataset of the Hungarian Tax Authority. Shares in the LEED are based on sums of firm-level and worker-level sample weights. A firm is foreign-owned if more than 50 percent of its equity is owned by foreign investors.

Figure 2: Changes in Selected Quantiles of the Real Log Wage Distribution, 1986-2008 (1986=100)



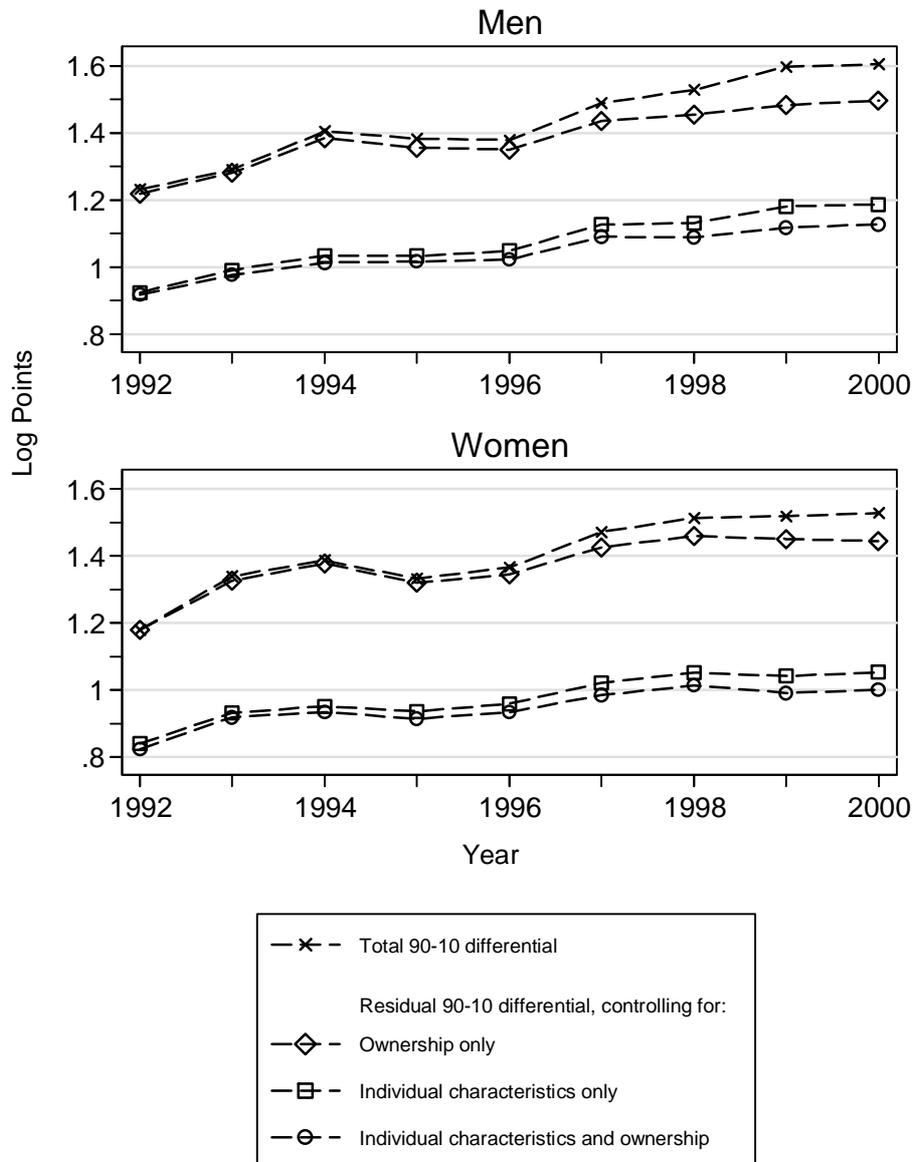
Notes: Indexed real changes in selected percentiles of the distribution of log monthly gross earnings. Each series is normalized separately by taking its 1986 value as the base period. Earnings measured in 2008 Hungarian forints. Years of large increases in the minimum wage: 2000 and 2001.

Figure 3: Evolution of Total, Between-Firm and Within-Firm Variance of Log Earnings, 1986-2008



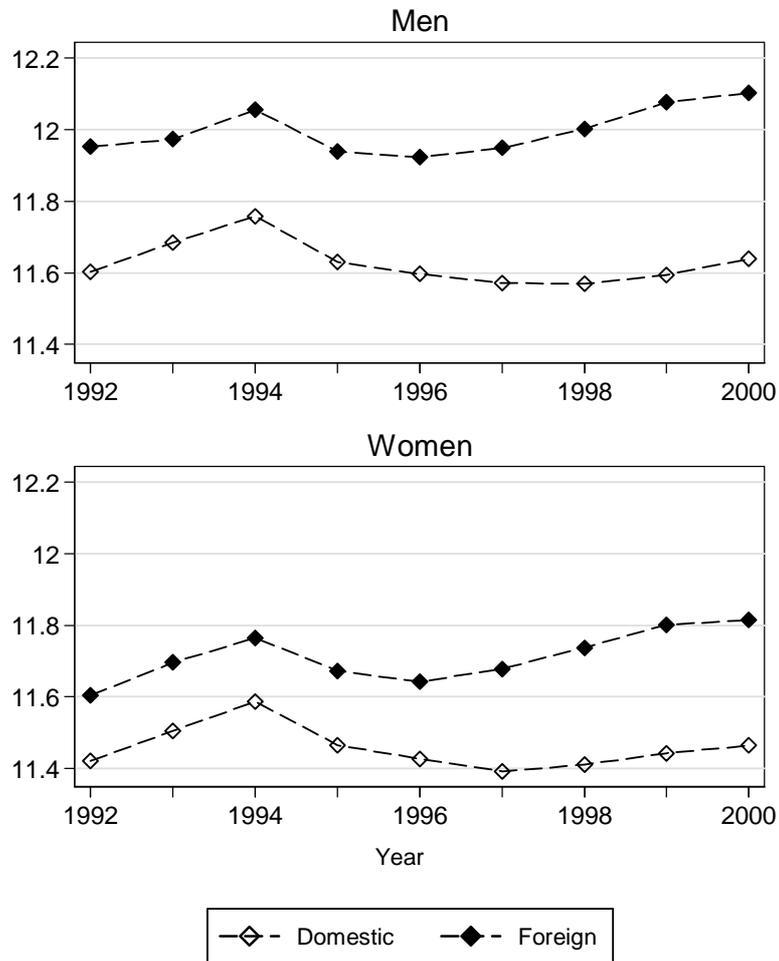
Notes: Results from a standard within-group/between-group variance decomposition of log real monthly gross earnings, performed by year. Groups of workers are defined as firms, and numbers of employees are used as group (firm) weights. Earnings measured in 2008 Hungarian forints.

Figure 4: Total and Residual Log 90-10 Wage Differentials, 1992-2000



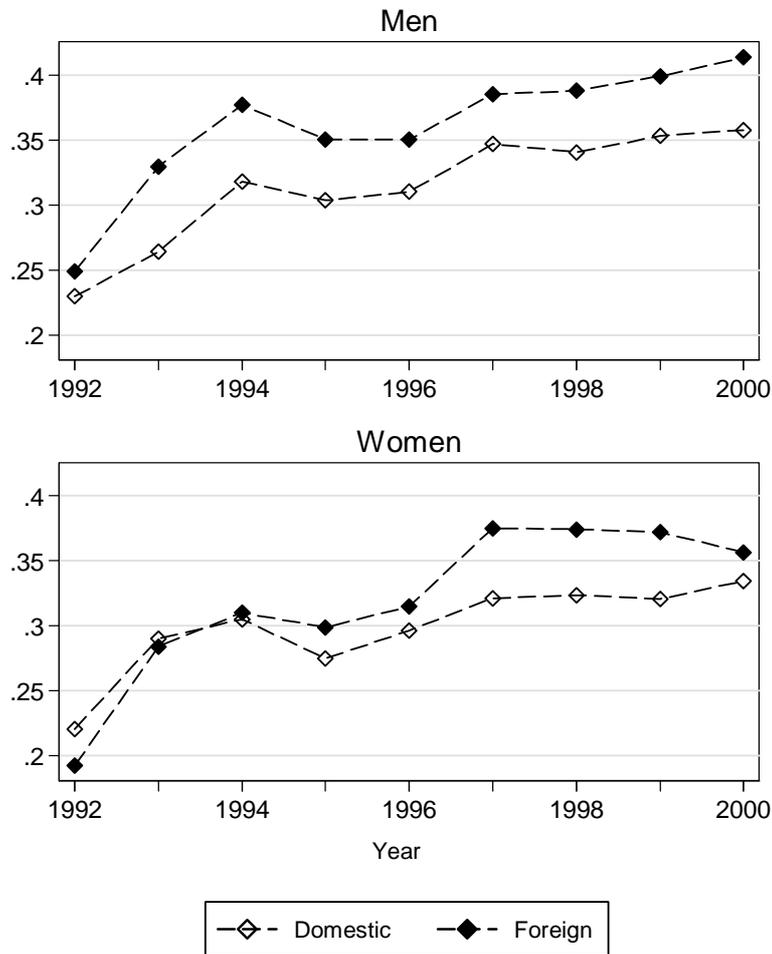
Notes: Residual interdecile differentials computed from the distribution of the residuals of yearly wage regressions. The dependent variable is log real monthly gross earnings, individual controls include highest degree of education, potential experience in levels, potential experience squared and full sets of occupational, industrial and regional dummies. Earnings measured in 2008 Hungarian forints. In the “Ownership only” regressions, only a constant and a lagged foreign ownership dummy are included. A firm is foreign-owned if more than 50 percent of its equity is owned by foreign investors.

Figure 5: Evolution of the Mean of Log Earnings by Ownership and Gender, 1992-2000



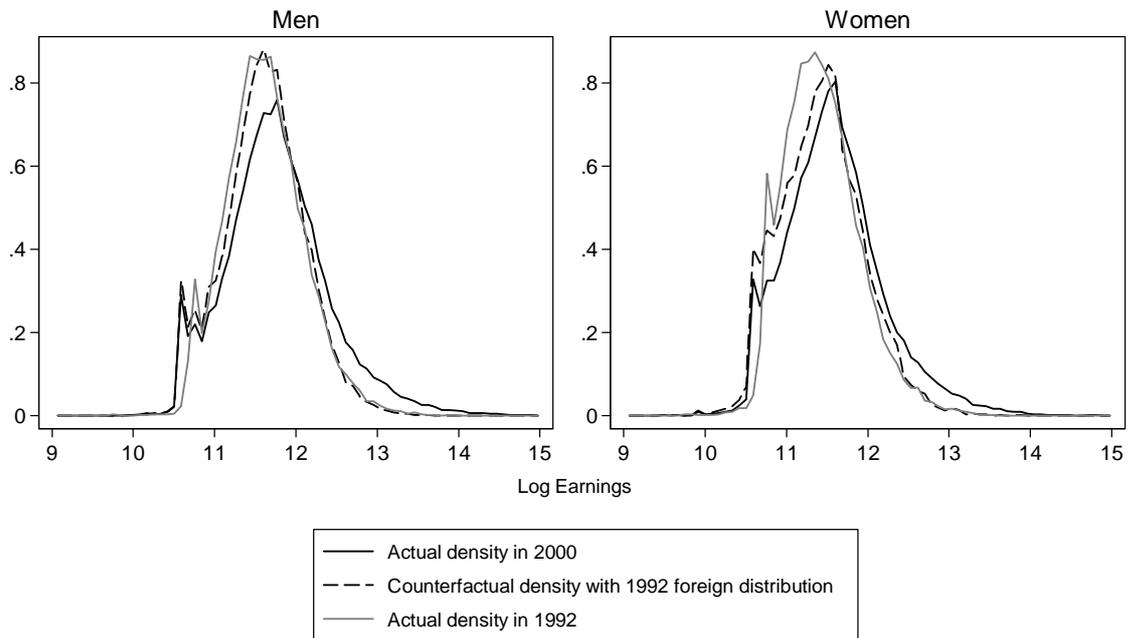
Notes: Unconditional means of log real monthly gross earnings computed separately for workers of domestic-owned and foreign-owned firms. Earnings measured in 2008 Hungarian forints. A firm is foreign-owned if more than 50 percent of its equity is owned by foreign investors.

Figure 6: Evolution of the Variance of Log Earnings by Ownership and Gender, 1992-2000



Notes: Unconditional variances of log real monthly gross earnings computed separately for workers of domestic-owned and foreign-owned firms. Earnings measured in 2008 Hungarian forints. A firm is foreign-owned if more than 50 percent of its equity is owned by foreign investors.

Figure 7: Composition Effect of Foreign Ownership on the Density of Log Earnings, 1992-2000

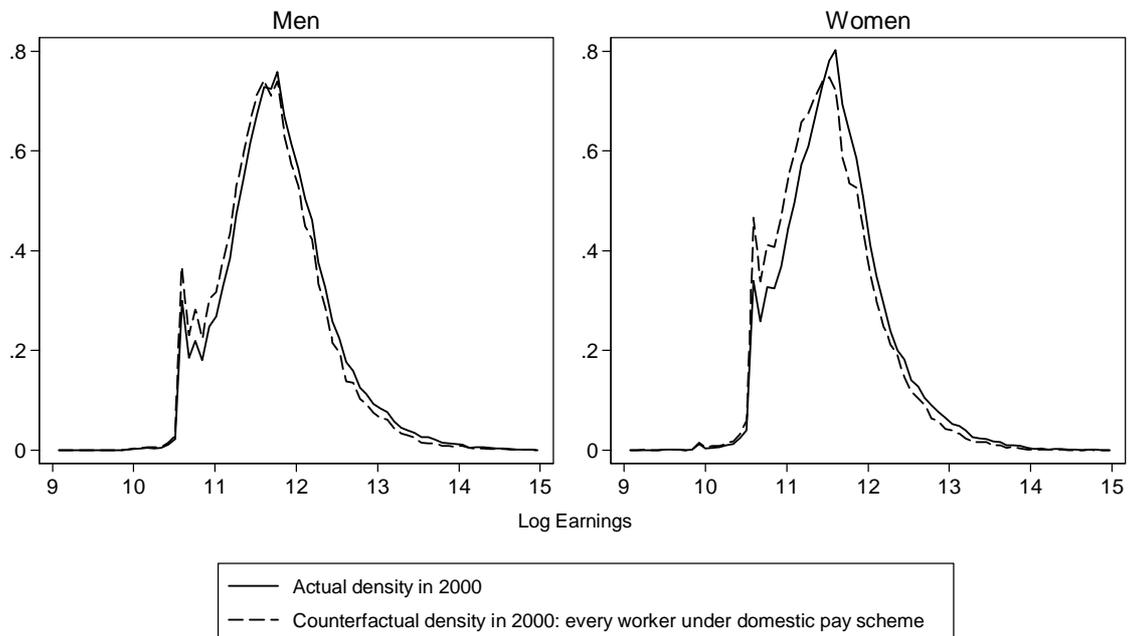


Notes: For the construction of the counterfactual density, the 2000 sample is reweighted to reflect the distribution of ownership in 1992, keeping the distribution of every other worker characteristic unchanged. Weights constructed by the reweighting method of DiNardo, Fortin and Lemieux (1996). See text for more details.

Figure 8a: Wage Structure Effect of Foreign Ownership on the Density of Log Earnings, 1992

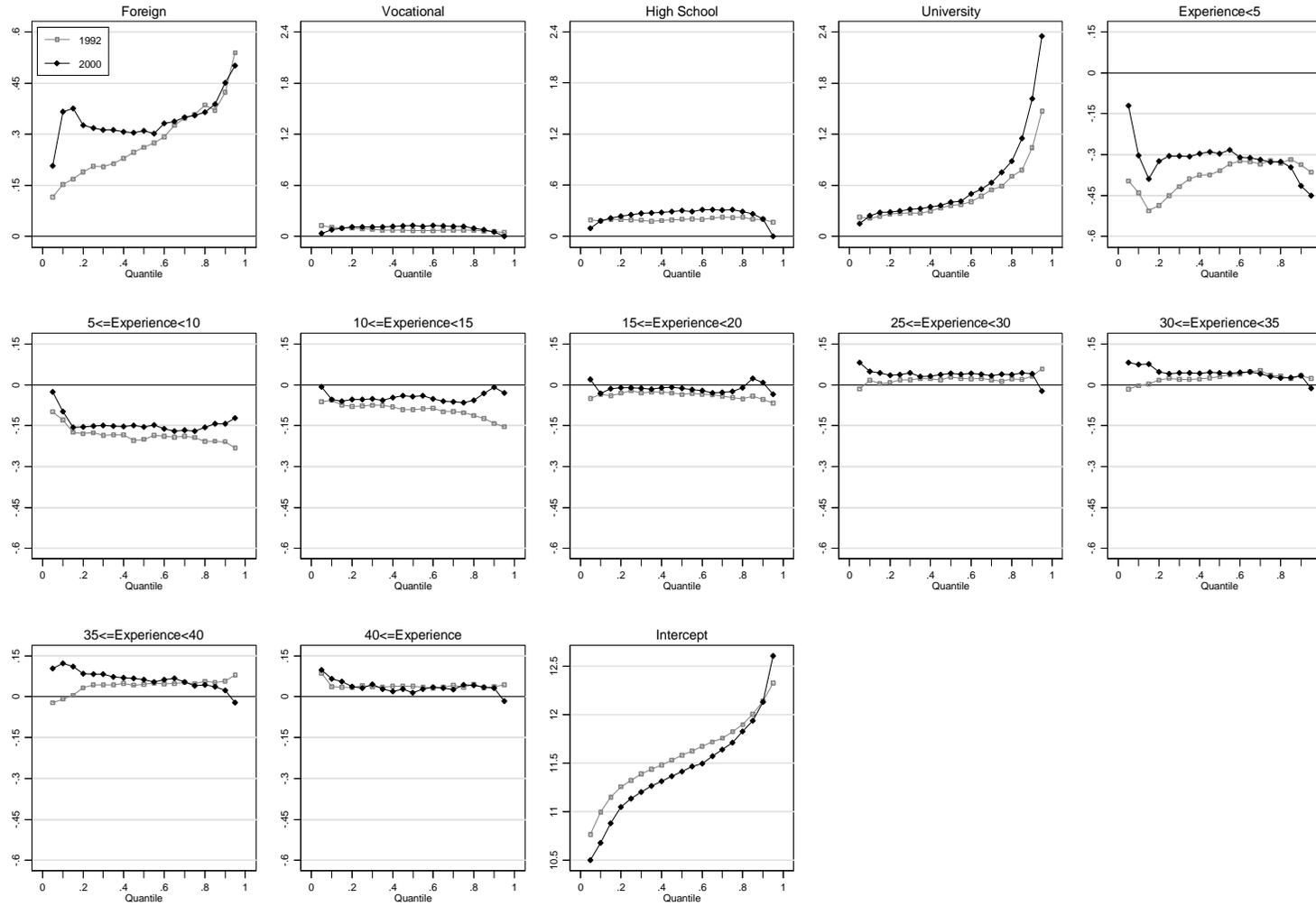


Figure 8b: Wage Structure Effect of Foreign Ownership on the Density of Log Earnings, 2000



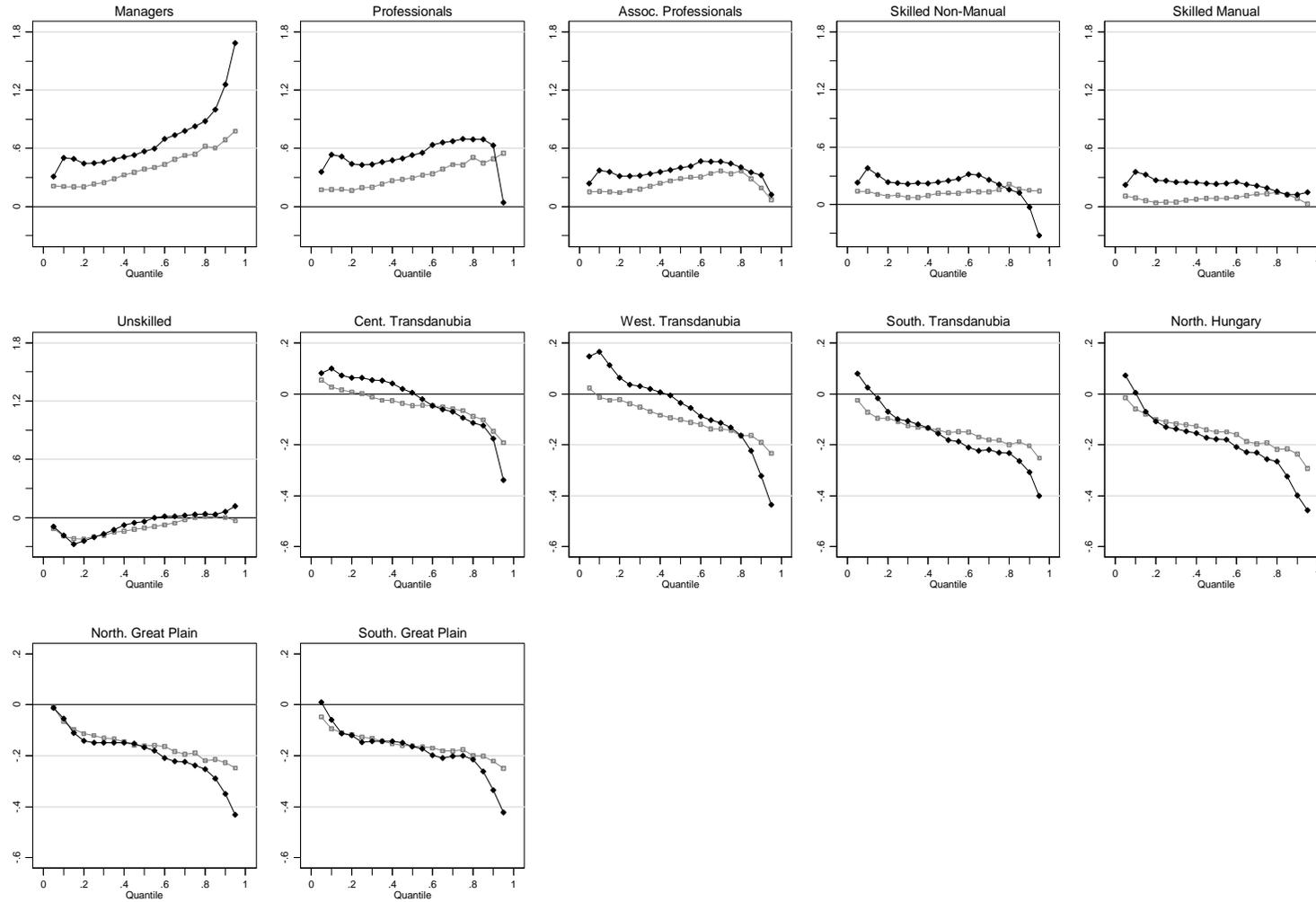
Notes: For the construction of the counterfactual density, the subsample of domestic workers is reweighted in each year to reflect the distribution of the total workforce, using the reweighting method of DiNardo and Lemieux (1997). See text for more details.

Figure 9a: Coefficients from Unconditional Quantile Regressions – Men



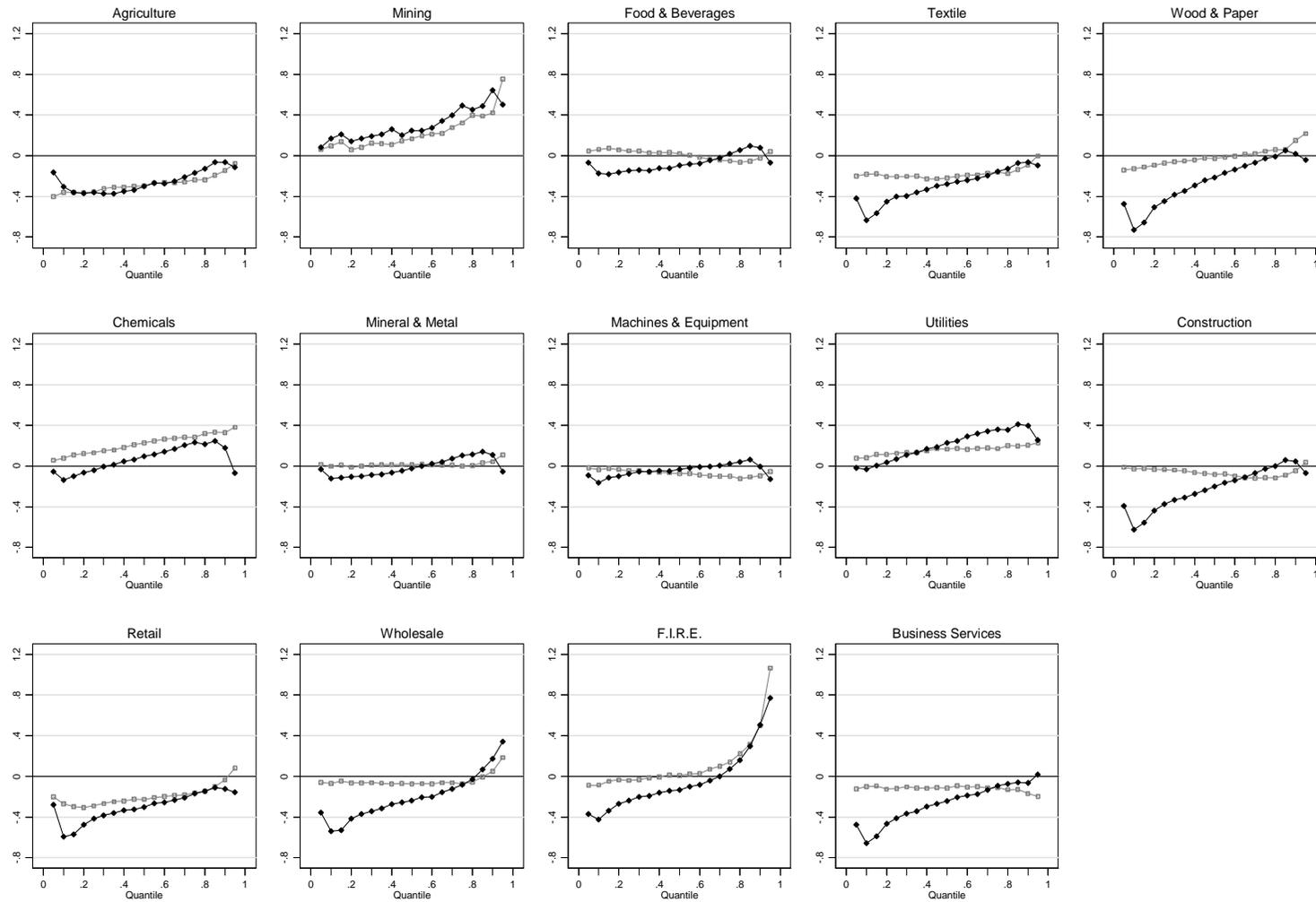
Notes: Omitted group: elementary education, more than 20 and less than 25 years of potential experience, services occupations, Central Hungary, other services industry.

Figure 9b: Coefficients from Unconditional Quantile Regressions – Men



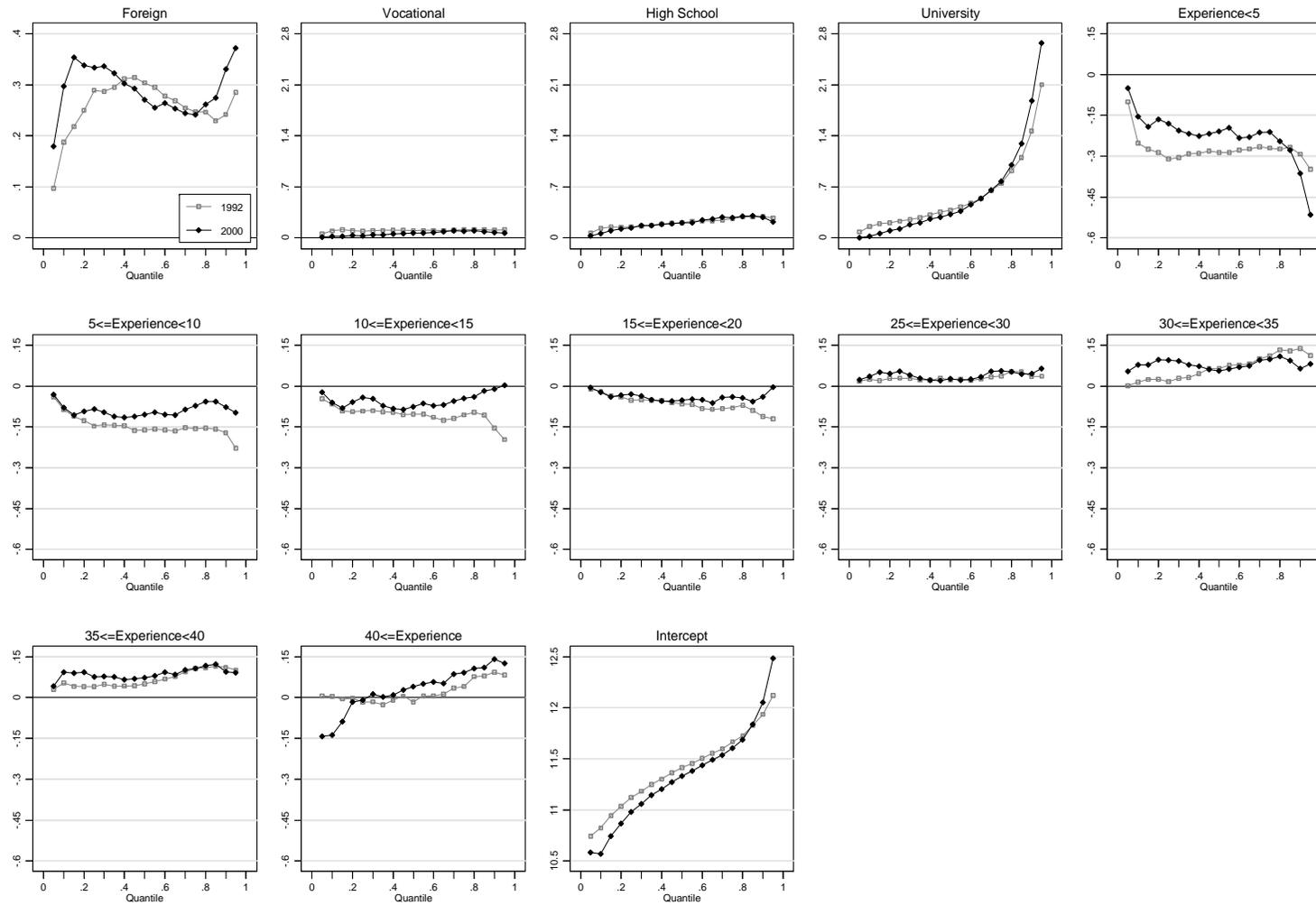
Notes: Omitted group: elementary education, more than 20 and less than 25 years of potential experience, services occupations, Central Hungary, other services industry.

Figure 9c: Coefficients from Unconditional Quantile Regressions – Men



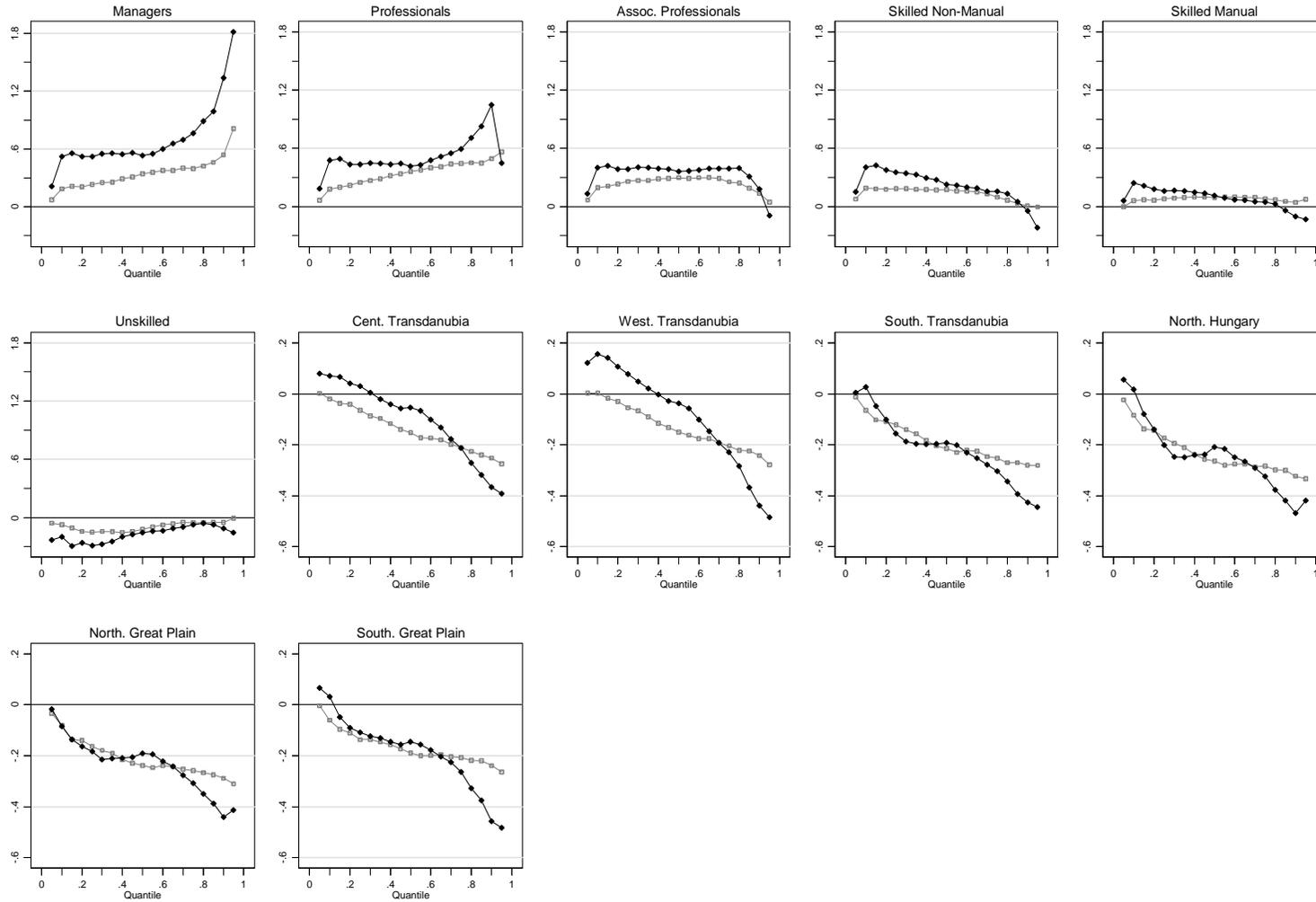
Notes: Omitted group: elementary education, more than 20 and less than 25 years of potential experience, services occupations, Central Hungary, other services industry.

Figure 10a: Coefficients from Unconditional Quantile Regressions – Women



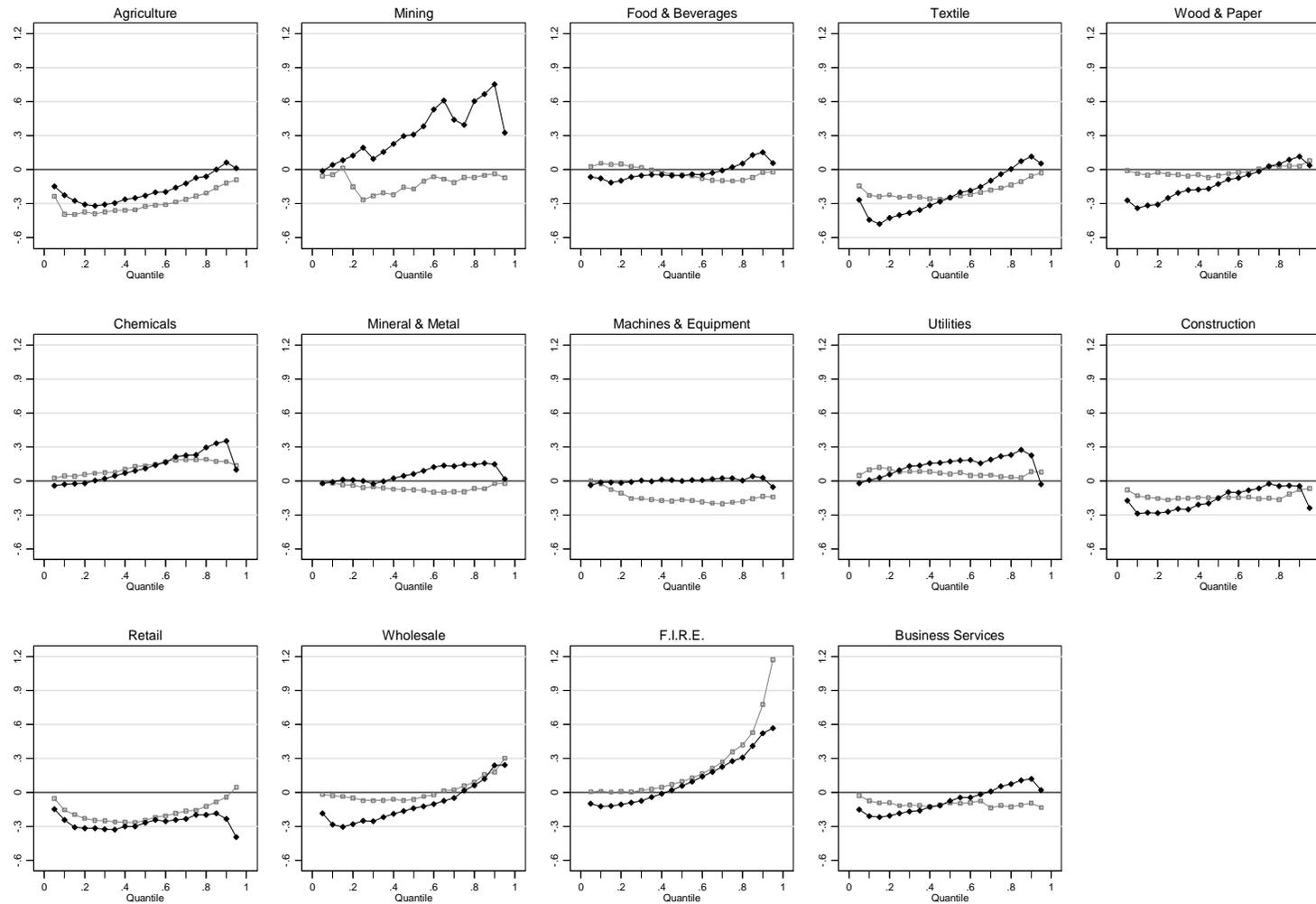
Notes: Omitted group: elementary education, more than 20 and less than 25 years of potential experience, services occupations, Central Hungary, other services industry.

Figure 10b: Coefficients from Unconditional Quantile Regressions – Women



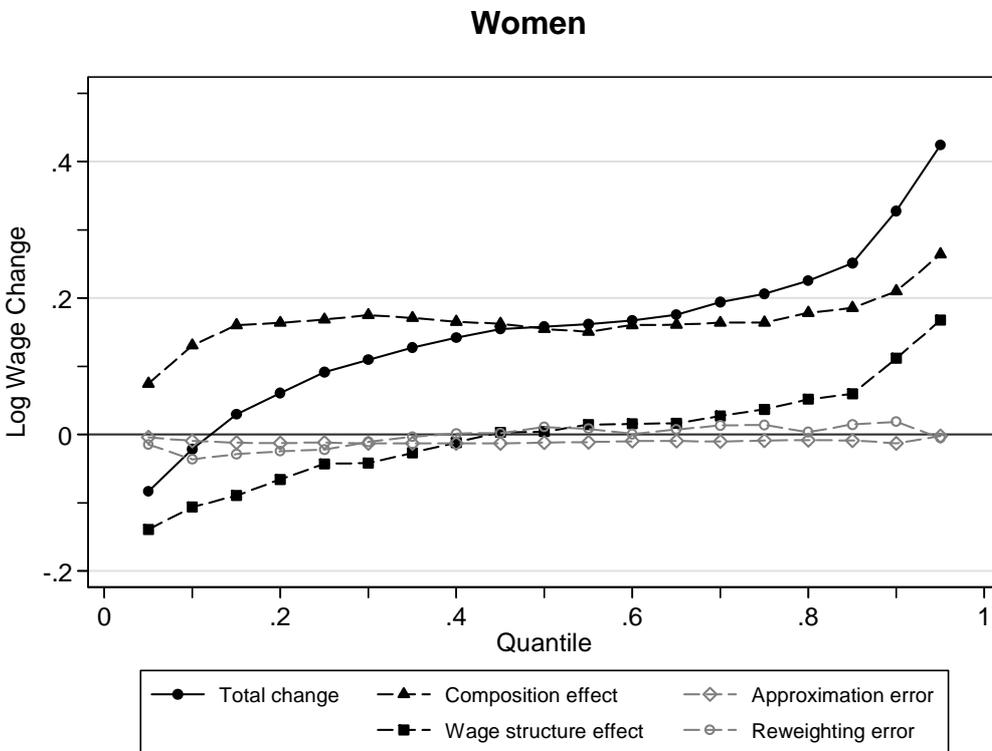
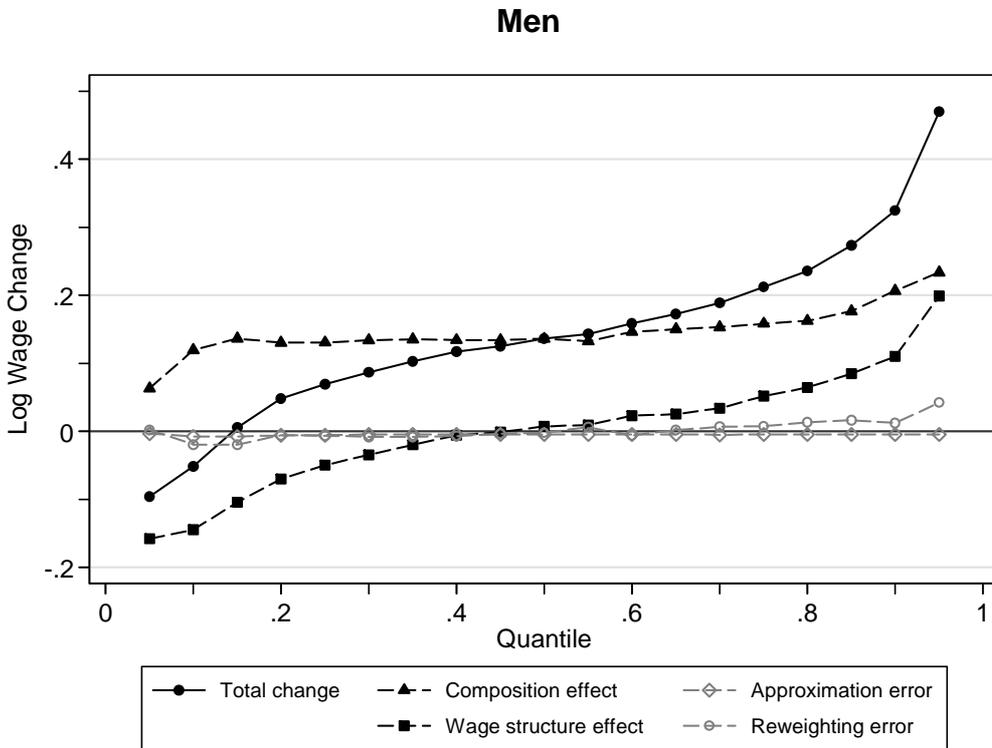
Notes: Omitted group: elementary education, more than 20 and less than 25 years of potential experience, services occupations, Central Hungary, other services industry.

Figure 10c: Coefficients from Unconditional Quantile Regressions – Women



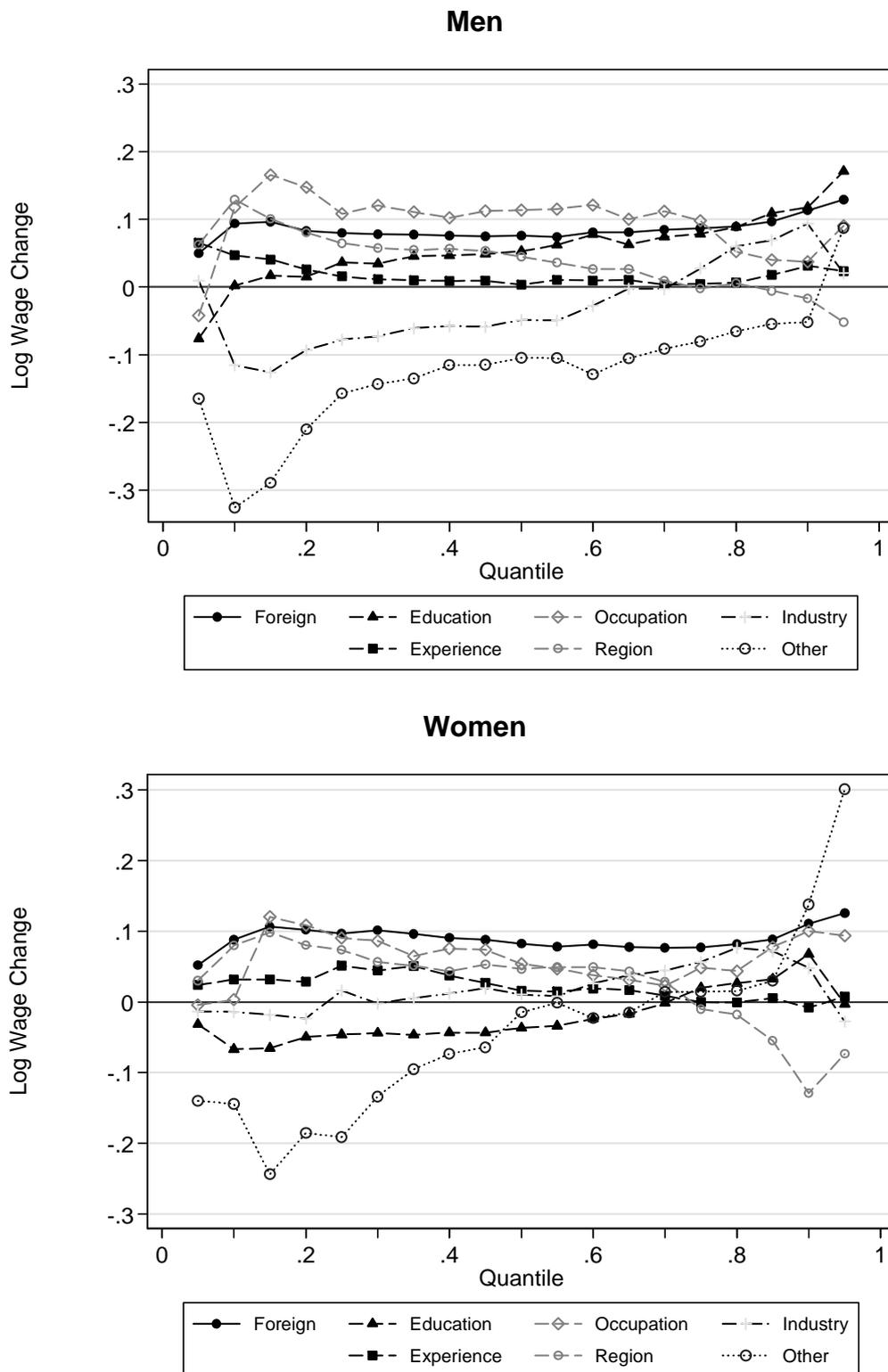
Notes: Omitted group: elementary education, more than 20 and less than 25 years of potential experience, services occupations, Central Hungary, other services industry.

Figure 11: Aggregate Decomposition of Total Wage Changes by Quantile (1992-2000)



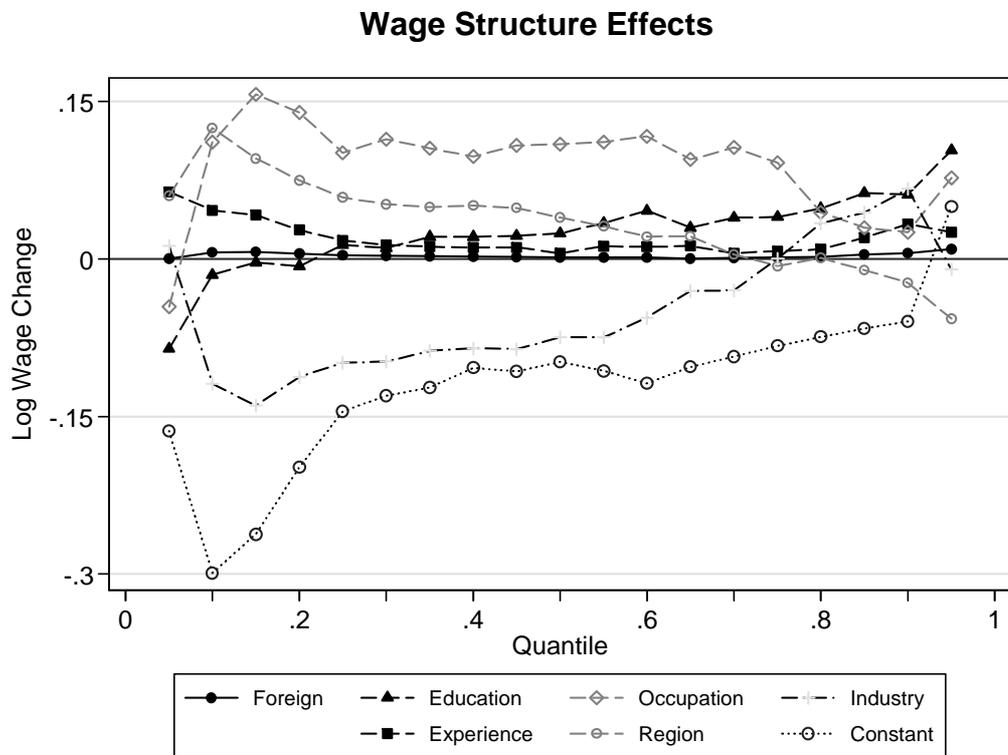
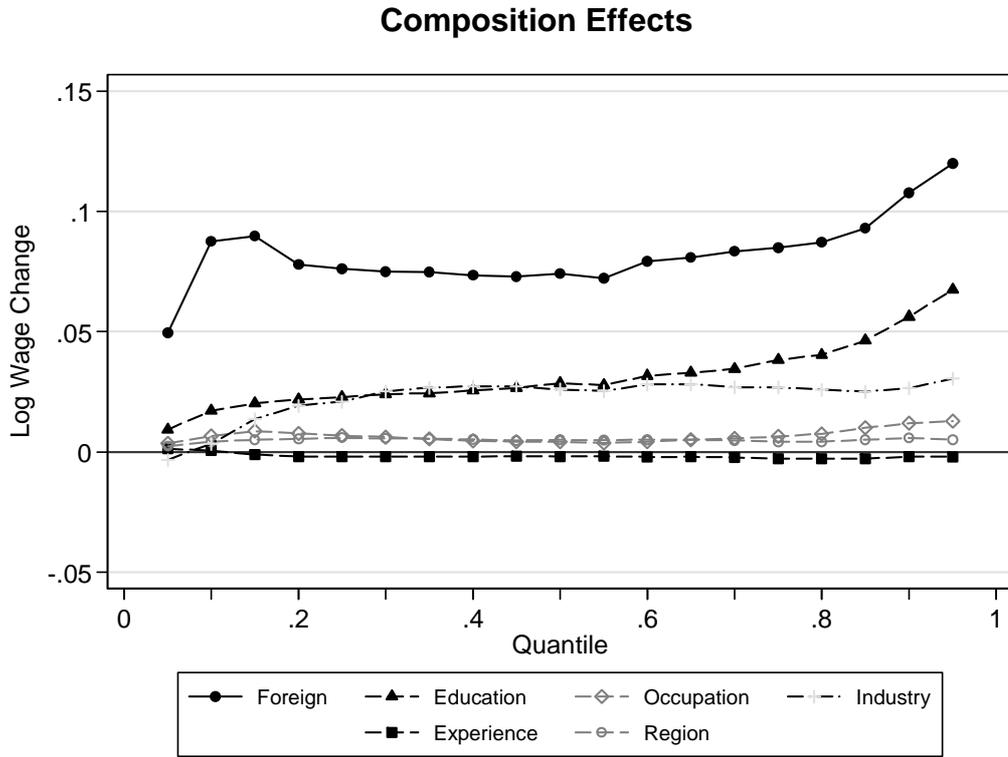
Notes: Based on a reweighted RIF-regression decomposition. Changes measured between 1992 and 2000, and a counterfactual outcome for which the distribution of worker characteristics in 2000 was reweighted to mimic the 1992 distribution.

Figure 12: Detailed Decomposition: Total Contributions of Worker Characteristics (1992-2000)



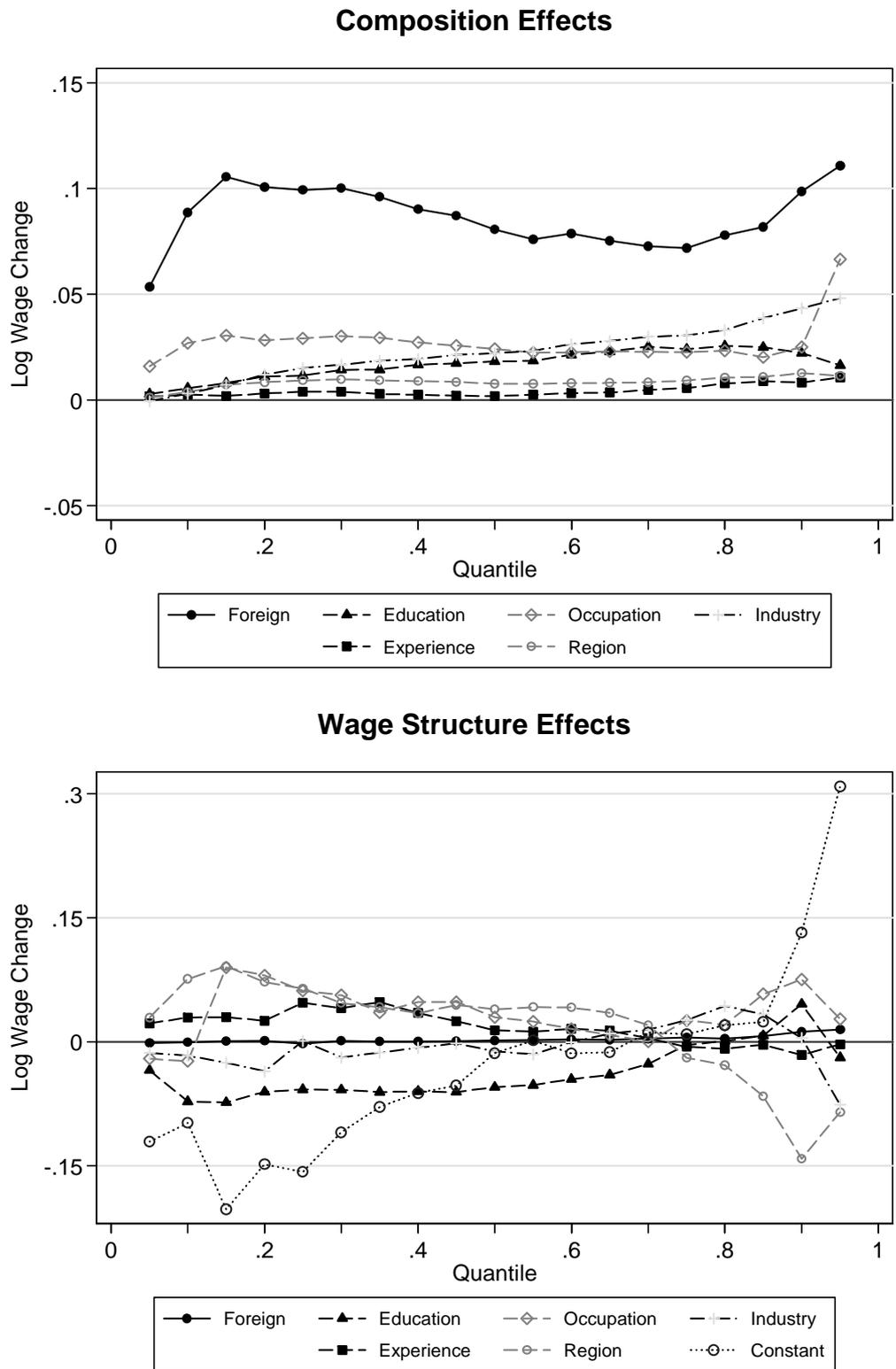
Notes: Based on a reweighted RIF-regression decomposition. Changes measured between 1992 and 2000 and a counterfactual outcome, for which the distribution of worker characteristics in 2000 was reweighted to mimic the 1992 distribution. "Other" includes the constant of the wage structure decomposition, and the approximation and specification errors.

Figure 13a: Detailed Composition and Wage Structure Effects - Men (1992-2000)



Notes: See Figure 11.

Figure 13b: Detailed Composition and Wage Structure Effects - Women (1992-2000)



Notes: See Figure 11.

Table 1: Sample Size by Year

Year	Unweighted	Weighted	
	Worker Observations	Employment in the Sample	Percent of Total Employment
1986	523,651	3,205.5	N.A.
1989	355,896	2,389.2	N.A.
1992	78,593	1,705.3	83.1
1993	81,875	1,327.4	74.8
1994	100,641	1,467.1	86.6
1995	102,634	1,482.7	92.1
1996	87,418	1,208.0	76.4
1997	85,451	1,157.6	73.5
1998	86,400	1,178.9	73.2
1999	84,319	1,119.7	70.3
2000	89,919	1,119.8	69.1
2001	86,759	1,109.1	68.6
2002	99,234	1,029.1	65.3
2003	96,143	954.1	61.0
2004	105,808	1,004.6	63.9
2005	113,058	1,033.9	66.7
2006	112,432	1,160.8	73.2
2007	104,788	971.5	62.4
2008	103,893	1,012.2	63.8
Total	2,498,912	N.A.	N.A.

Note: Employment in the Sample (in thousands) = sum of workers employed by firms in the LEED. Percent of Total Employment = sample employment divided by total employment in the dataset of the Hungarian Tax Authority (HTA). Only firms with more than 20 employees included. The HTA dataset contains virtually every double-entry book-keeping company in the business sector, except for years before 1992 when it includes only a sample of firms.

Table 2a: Variance of Log Earnings Within- and Between Ownership Groups (Variance Decomposition of Levels)

	Total Variance	Within-Group Variance	Between-Group Variance	Group-Level Variances		Employment Shares	
				Domestic	Foreign	Domestic	Foreign
Men							
1992	0.235	0.230	0.005	0.230	0.249	0.960	0.040
2000	0.415	0.373	0.043	0.358	0.414	0.727	0.273
Women							
1992	0.221	0.219	0.002	0.221	0.193	0.943	0.057
2000	0.370	0.342	0.028	0.335	0.357	0.647	0.353

Notes: Results from a standard within-group/between-group variance decomposition performed by year, where groups of workers are defined as workers of domestic and foreign firms, and numbers of employees are used as group weights.

Table 2b: Changes in the Variance of Log Earnings Within- and Between Ownership Groups (Variance Decomposition of Changes)

	Total Change in Variance	Within-Group Change		Between-Group Change		Change in Domestic Variance	Change in Foreign Variance
		Change in Variance	Composition Effect	Change in Variance	Composition Effect		
Men							
1992-2000	0.180	0.129	0.013	0.015	0.023	0.128	0.165
Women							
1992-2000	0.149	0.117	0.007	0.016	0.011	0.114	0.164

Notes: Changes in total variance decomposed according to the decomposition method in Table 2a.

Appendix Table 1: Descriptive Statistics by Ownership Type

	1992		2000	
Foreign Employment Share (%)	4.6		30.5	
	Domestic	Foreign	Domestic	Foreign
Monthly Earnings	116.1 (71.1)	152.4 (104.7)	131.5 (134.7)	202.4 (225.2)
Female (%)	37.4	46.7	37.0	46.0
Education (%)				
<i>Elementary</i>	32.9	33.8	23.6	19.6
<i>Vocational</i>	32.7	33.5	38.4	33.7
<i>High school</i>	27.1	23.5	29.9	32.7
<i>University</i>	7.3	9.2	8.2	14.0
Experience	22.1 (10.6)	20.4 (10.5)	23.1 (10.9)	19.3 (10.9)
Occupation (%)				
<i>Elementary Occupations</i>	11.1	10.8	9.6	5.4
<i>Skilled Manual Workers</i>	48.3	58.2	50.5	53.3
<i>Service Workers</i>	9.2	5.0	10.9	7.2
<i>Clerks</i>	6.8	5.4	5.9	6.0
<i>Associate Professionals</i>	12.7	9.6	12.1	14.5
<i>Professionals</i>	6.2	6.9	2.9	6.2
<i>Managers</i>	5.7	4.0	8.2	7.3
Industry (%)				
<i>Agriculture</i>	18.3	0.2	12.3	0.6
<i>Mining</i>	0.2	0.0	0.3	0.0
<i>Food&Beverages</i>	6.2	11.1	6.5	7.6
<i>Textile</i>	5.4	12.3	6.8	9.8
<i>Wood&Paper</i>	2.6	2.6	3.1	2.5
<i>Chemicals</i>	4.8	3.7	2.7	9.5
<i>Minerals&Water</i>	5.3	4.9	6.7	7.5
<i>Machines&Equipment</i>	8.8	43.1	9.8	26.2
<i>Utilities</i>	3.0	0.0	2.8	5.1
<i>Construction</i>	6.1	8.8	6.3	1.8
<i>Retail Trade</i>	9.5	7.2	7.3	7.4
<i>Wholesale Trade</i>	4.0	4.0	4.0	5.2
<i>F.I.R.E.</i>	1.5	0.1	4.5	5.9
<i>Business Services</i>	2.6	1.1	4.9	3.6
<i>Other Services</i>	21.7	0.8	22.0	7.3
N	74,724	3,869	59,987	29,932

Notes: Weighted unconditional means and standard deviations. Earnings measured in thousands of 2008 HUF, deflated by CPI. Female, education, new hire and occupation measured as percentages of total workforce by ownership type. Standard deviations in parentheses. The definition of occupations follows ISCO-88 where Elementary Occupations, Service Workers, Clerks, Associate Professionals, Professionals and Managers coincide with the corresponding major groups; while Skilled Manual Workers cover Skilled agricultural and fishery workers, Craft and related trades workers and Plant and machine operators and assemblers.

**Appendix Table 2: Estimated Coefficients of Foreign Ownership in
Unconditional Quantile Regressions**

	Men		Women	
	1992	2000	1992	2000
1 st Decile	0.152** (0.021)	0.366** (0.033)	0.188** (0.021)	0.297** (0.030)
2 nd Decile	0.190** (0.023)	0.326** (0.022)	0.250** (0.029)	0.338** (0.028)
3 rd Decile	0.204** (0.028)	0.313** (0.018)	0.287** (0.036)	0.336** (0.026)
4 th Decile	0.229** (0.035)	0.307** (0.018)	0.312** (0.046)	0.303** (0.022)
Median	0.262** (0.041)	0.310** (0.020)	0.304** (0.045)	0.271** (0.020)
6 th Decile	0.292** (0.044)	0.331** (0.025)	0.277** (0.047)	0.264** (0.023)
7 th Decile	0.347** (0.051)	0.349** (0.029)	0.255** (0.036)	0.244** (0.027)
8 th Decile	0.386** (0.058)	0.364** (0.032)	0.246** (0.032)	0.261** (0.033)
9 th Decile	0.424** (0.057)	0.451** (0.045)	0.242** (0.035)	0.331** (0.039)
N	44,072	50,495	31,887	37,235

Notes: The table shows estimated coefficients on the foreign ownership dummy from RIF regressions. Other controls include education, experience, region, industry and occupation. See text for more details. Standard errors in parentheses. ** = significant at 0.01; * = significant at 0.05

Appendix Table 3: Contributions of FDI to Changes in Log Wage Differentials, 1992-2000

	90-10	90-50	50-10
Men			
Total Change	0.376	0.187	0.189
FDI Composition Effect	0.021	0.034	-0.013
FDI Wage Structure Effect	-0.001	0.003	-0.004
Women			
Total Change	0.350	0.170	0.180
FDI Composition Effect	0.010	0.018	-0.008
FDI Wage Structure Effect	0.013	0.001	0.003

Notes: Computed from the results of RIF decompositions presented in Figures 11-13. Changes measured in log points.