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Labour Demand with Heterogeneous Labour Inputs after the Transition in Hungary, 1992–1999 – and the Potential Consequences of the Increase of Minimum Wage in 2001 and 2002

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LABOUR DEMAND WITH HETEROGENEOUS LABOUR INPUTS AFTER THE TRANSITION IN HUNGARY, 1992–1999 – AND THE POTENTIAL CONSEQUENCES OF THE INCREASE OF MINIMUM WAGE IN 2001 AND 2002

by

GÁBOR KERTESI and **JÁNOS KÖLLŐ**

Abstract

The paper analyses changes in the demand for unskilled, young skilled, and older skilled workers during the post-communist transition in Hungary. Systems of cost share equations derived from the translog cost function are estimated for cross-sections of large firms observed in the period 1992-99. Following the 'transformational recession' the own-price elasticities of labour and capital were stabilized at levels observed in several developed market economies. Unskilled and skilled labour are estimated to be p-complements, and younger and older skilled workers p-substitutes. Capital and labour appear to be p-substitutes with unskilled labour having the highest elasticity of substitution. Further results hint at the existence of non-negligible scale effects and the non-neutrality of technical change. The estimated wage elasticities give us the opportunity to evaluate consequences of some governmental policies. As minimum wage was doubled in nominal terms between 1999 and 2002 in Hungary it was evident to apply these results to this highly relevant issue. In the second part of the paper we try to evaluate the potential demand consequences of this. Based on the earnings distributions of the Wage Survey of 1999 (a large individual level data set) we make several predictions concerning these consequences.

A VÁLLALATI MUNKAERŐ-KERESLET MAGYARORSZÁGON 1992 ÉS 1999 KÖZÖTT ÉS A 2001/2002. ÉVI MINIMÁLBÉR-EMELÉSEK VÁRHATÓ FOGLALKOZTATÁSI KÖVETKEZMÉNYEI

(Heterogén munkaerő-keresleti modell)

KERTESI GÁBOR – KÖLLŐ JÁNOS

Összefoglaló

A gazdaságban alkalmazott munkaráfordításokat különböző típusokra bontottuk. Megkülönböztettük azokat a munkafajtákat, melyeket alacsony iskolázottságú ("képzetlen"), illetve azokat, amelyeket legalább középiskolai végzettségű ("képzett") munkaerő lát el. Az utóbbi kategóriát további két csoportra bontottuk aszerint, hogy a jelenlegi iskolázott munkaerő tudását milyen időszakban – tipikusan a rendszerváltás előtti évtizedekben vagy a nyolcvanas-kilencvenes években – szerezte meg. Összességében így három munkaerőcsoportot különböztettünk meg: képzetlen, fiatal-képzett és idős-képzett munkaerőt, melyeket jellegzetesen különböző termelési tényezőknek tekintettünk. A tanulmány vállalati szintű négytényezős munkaerő-keresleti függvényeket becsül e háromféle munkaráfordításra és a tőkére mint negyedik erőforrásra alapozva, a 300-nál több alkalmazottat foglalkoztató magyarországi vállalatok mintáján, keresztmetszetben, az 1992 és 1999 közti időszak valamennyi évére. A becslés alapjául transzlog költségfüggvényből származtatott költségarány-egyenletek szimultán becslése szolgált. Az alábbi eredményekre jutottunk. A különböző munkafajták és a tőke saját-árrugalmasságai a stabil piac-gazdaságokban megszokott értéket veszik fel. Az iskolázatlan munkaerő kereslete különösen érzékeny a bérváltozásra. Ami a kereszt-árrugalmasságokat illeti: az iskolázatlan és az iskolázott munkaerő egymás kiegészítői, függetlenül attól, hogy fiatal vagy idős iskolázott munkaerőről van szó. Mindhárom munkafajta helyettesítő viszonyban áll a tőkével, a képzetlen munkaerő azonban különösen könnyen váltható ki tárgyi eszközökkel, amennyiben a bére emelkedik. A becslési eredmények nem elhanyagolható mérethatásra, illetve a külföldi és belföldi többségi tulajdonlással együttjáró nem semleges hatékonyságkülönbségek jelenlétére is utaltak. A becsült elaszticitások felhasználhatók arra is, hogy a munkapiac működésébe való egyes állami beavatkozások várható hatásait felmérjük. A tanulmány a 2001. januári, illetve a 2002. januári drasztikus minimálbér-emelések várható foglalkoztatási következményeit a munkaerő-keresleti modell saját-bérrugalmasságaira és az OMMK 1999. évi bértarifafelvételének egyéni béreloszlásaira támaszkodva becsülte meg. Számításaink a minimálbér-emelésből fakadóan komoly mértékű foglalkoztatáscsökkenési hatást jeleznek előre a képzetlen munkaerő kategóriáján belül országosan is, az ország elmaradottabb (alacsony átlagos iskolázottságú) vidékein pedig különösen.

1. INTRODUCTION

Transition to the market economy decreased the demand for unskilled labour in Central and East European countries and (according to the majority of the estimates) devalued the skills of older workers. This paper would like to contribute to a better understanding on this process by studying the degree of substitution and/or complementarity among skill groups and generations in the Hungarian large-firm sector.

Most of what is known about the revaluation of human capital in Central and Eastern Europe comes from estimates of Mincer-type earnings functions. The adverse implications of systemic change for unskilled workers are well known and supported by robust evidence from all countries of the region. Most studies observed declining returns to experience, too, but in this case the empirical evidence is far from being unambiguous. Rutkowski (1997) and Puhani (1997) presented evidence of falling returns in *Poland* 1987–92 and 1992–95, respectively, but not later. In the former *Czechoslovakia* Vecernik 1995, Sakova 1998, Flanagan 1995 and Chase 1997 observed declining returns in early stages of the transition but a study using retrospective wage data by Munich, Svejnar and Terrell (2000) detected no change in the experience-wage profile between 1989 and 1996.

Several papers on the *former GDR* (Steiner and Bellmann (1997), Burda and Schmidt (1997), Krueger and Pischke, (1992)) suggested that the returns to experience fell after re-unification but no decline was observed by Steiner and Wagner (1997) in their female sub-sample. In contrast to other studies Franz and Steiner (1999) estimated completely flat experience-wage profiles for men both before and after the unification, and falling returns for women. In *Hungary* Kertesi and Köllő (2001) found that the general devaluation of experience and the appreciation of new skills (manifesting itself in particularly fast-rising returns to education in young cohorts) continued at least until 1999 but the relative wages of older skilled workers failed to increase after 1992.

What happened and why is not easy to tell on the basis of the reduced-form Mincer-type earnings function estimates since they capture the combined effect of a variety of demand and supply side mechanisms. This paper would like to make a step further by explicitly addressing one of the underlying mechanisms – changes in firms' demand for education-based and age-specific skills. A translog cost function is used to estimate the relevant elasticities for cross-section samples of large Hungarian firms observed between 1992 and 1999. The labour demand model is based on clear

(though not necessarily correct) assumptions that provide handles for the interpretation and discussion, more so than do the earnings functions.

Beside the wish to 'open the black box' the research is motivated by several practical questions concerning the future rather than the past of the labour market in a country where the social divide between skilled versus unskilled and young versus old workers is apparent. Just to mention some highly relevant issues: (i) Can we expect a further decline of demand for unskilled and older skilled workers or the process is likely to halt due to the falling relative wages of these groups and/or by the recovery of output? (ii) In view of the substantial expansion of general secondary and higher education during the transition period can we expect a 'crowding out' effect – the substitution of young and skilled labour for unskilled and old-and-skilled labour on a massive scale – or the likely outcome will rather be a deterioration of the career prospects of the younger generation? (iii) How labour demand is affected by the change in the minimum wage level? As minimum wage was doubled in nominal terms between 1999 and 2002 in Hungary this is an issue of great practical importance now. The second part of this paper is devoted to the evaluation of the potential employment consequences of this huge rise of minimum wage. The discussion of these issues presupposes the knowledge of the own-wage, cross-wage and output elasticities of demand for different types of skills. This paper would like to add some first, crude empirical results to the discussion, which, in lack of research on the issue, has been based on conjectures and analogies with the Western experience.

2. A MODEL OF LABOUR DEMAND WITH HETEROGENOUS LABOUR INPUTS

2.1 Analytical framework

The interest in multi-factor demand systems has been traditionally connected with the issue of social equity and equality-enhancing employment policies. Typical examples are Borjas (1983) analysing the competition for jobs between blacks, Hispanics and whites on the US labour market, or Grant and Hamermesh (1981) studying substitution between various social groups (white women, youths and others).

Another source of interest becoming increasingly important in the last two decades has been the impact of technological change on the demand for skills via complementarity or substitution with non-labour inputs and among various types of skills. Among the pioneers in this field were Grili-

ches (1969), Berndt and Christensen (1973), Freeman (1979), Freeman and Medoff (1982). Skill-biased technological change combined with the rigidity of wages gave rise to growing interest for the topic in Europe. (See recent papers by Steiner and Wagner (1997), Steiner and Mohr (1998), Falk and Koebel (2000) on contemporary Germany, for instance).

The approach of this paper is closest to Freeman (1979) in that the generational divide is treated as a central issue and demand is analysed by means of a translog specification of the production and cost functions. Since we look at enterprises free to set factor quantities but not factor prices we shall prefer the cost function as a tool for analysis.

To start with, we assume that technologies can be described with a translog production function. Output (Y) is produced using capital and three types of labour distinguished by education and experience. Factor quantities and factor prices are denoted with X_i and p_i $(i = 1,2,3,4)$ respectively. Corresponding to the production function exists a minimum cost (C*) function:

$$\ln C^* = v_0 + v_y \ln Y + \sum_{i=1}^{4} v_i \ln p_i + \frac{1}{2}\sum_{i=1}^{4}\sum_{j=1}^{4} \gamma_{ij} \ln p_i \ln p_j + \sum_{i=1}^{4} \gamma_{iy} \ln p_i \ln Y + R_m \quad (1)$$

As shown in Binswanger (1974), Sato and Koizumi (1975), Hamermesh (1991, 1993) and elsewhere starting from the cost function, applying Shephard's lemma, and neglecting the remainder one can derive the following system of optimal cost share equations:

$$\frac{\partial \ln C^*}{\partial \ln p_i} = \frac{p_i X_i}{\sum_i p_i X_i} = s_i = v_i + \sum_{j=1}^{4} \gamma_{ij} \ln p_j + \gamma_{iy} \ln Y, \qquad i = 1,2,3,4 \quad (2)$$

The equality of the cross-derivatives and the homogeneity of degree one of the cost function in prices implies the following within and cross-equation constraints:

$$\gamma_{ij} = \gamma_{ji} \quad (i \neq j) \quad (3a)$$

$$\sum_{i=1}^{4} \gamma_{ij} = 0, \; j = 1,2,3,4 \quad (3b)$$

$$\sum_{i=1}^{4} \gamma_{iy} = 0 \quad (3c)$$

The own-price elasticities (ε_{ii}) and the cross-price or partial substitution elasticities (σ_{ij}) can be computed using the estimated parameters and observed cost shares:

$$\varepsilon_{ii} = \frac{\gamma_{ii} + s_i^2 - s_i}{s_i^2} \tag{4a}$$

$$\sigma_{ij} = \frac{\gamma_{ij} + s_i s_j}{s_i s_j} \tag{4b}$$

Models similar to (2)–(3) could be derived following similar logic from the CES and the generalized Leontief production or cost functions. If factor quantities were assumed to be exogenous instead of factor prices elasticities of complementarity (close relatives to the elasticities of substitution) could be derived easily from the production function.[1]

The empirical estimates of a model with four factors of production can be obtained from a system of three equations. Being linearly dependent on the remaining three equations the fourth one can be dropped from the model that thus reduces to:

$$\begin{aligned} s_{1n} &= v_1 + \sum_{j=1}^{3} \gamma_{1j} \ln p_{jn} + \gamma_{1y} \ln Y_n + \gamma_{1F} F_n + \xi_{1n} \\ s_{2n} &= v_2 + \sum_{j=1}^{3} \gamma_{2j} \ln p_{jn} + \gamma_{2y} \ln Y_n + \gamma_{2F} F_n + \xi_{2n} \\ s_{3n} &= v_3 + \sum_{j=1}^{3} \gamma_{3j} \ln p_{jn} + \gamma_{3y} \ln Y_n + \gamma_{3F} F_n + \xi_{3n} \end{aligned} \tag{5}$$

with $j = 1,2,3$ standing for the three groups of labour, $n = 1,2,...,N$ standing for the observed firms, and *F* denoting variables which control for non-neutral efficiency differences. (These will be introduced later). The model is subject to the constraints:

$$\gamma_{12} = \gamma_{21}, \qquad \gamma_{13} = \gamma_{31}, \qquad \gamma_{23} = \gamma_{32}. \tag{6}$$

The parameters for the capital share equation can be obtained from the constraints (3b) and (3c)[2] and the elasticities can be computed according to (4a) and (4b). The error terms are assumed to be randomly distributed with zero expected value, potentially correlated.

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[1] For deriving elasticities of substitution from the production function or elasticities of complementarity from the cost function the bordered Hessian of the estimates should be inverted. In practice, the exogeneity assumptions decide what type of elasticities will be computed.

[2] We have an additional cross-equation constraint for the parameters of the *F* variable: $\sum \gamma_{iF} = 0$

Though the dependent variables are the same in all equations the OLS estimator would still be inefficient because of the restrictions imposed in (6). Therefore Zellner's seemingly unrelated regression or GMM should be applied for a simultaneous estimation of the system. We shall apply the first.

A convenient alternative to estimating (5)–(6) would be distinguishing between only two types of labour. This would simplify the analysis in that the system would collapse to a single equation with a 'skills ratio' on the left hand and a 'wage ratio' (plus demand shift variables) on the right hand. This path is followed by Steiner and Mohr (1998) for instance who define skills ratios for industry-experience cells and regress them on the respective wage ratios and a trend variable using longitudinal data. Though the virtues of this type of analysis are evident the underlying assumptions seem too strong for an economy in transition. The longitudinal skills ratio model assumes that the relative efficiency of skilled versus unskilled labour is equal across the units of observations and changes at a constant rate over time. Since we can reasonably assume that education yields higher productivity returns in the case of young workers the first assumption (equal relative productivity of skilled versus unskilled labour in the industry-experience cells) is unlikely to hold. Second, as will become clear, there are severe structural breaks in the data calling into question the second assumption.

A static labour demand model should cope with at least three risks arising from its restrictive assumptions. Most importantly, being derived from equilibrium conditions the static model assumes that the observed cost shares represent optimal choice under cost minimization or, at least, are randomly dispersed around the optimum. Estimates relating to the early stages of the transition may potentially suffer from the invalidity of this assumption. A formal model of factor demand during an economic turmoil like the 'transformational recession' would be useful in relaxing the assumptions but the available theorems (Kornai (1992) in particular) seem to us too complex for being incorporated into the framework used here.

A further problem of static factor demand models is connected with the assumption that firms are located on the same production isoquant. In case some firms are more efficient than others and the efficiency differentials are systematically related to the firms' mix of factors of production (that is, efficiency differentials are non-neutral) the model produces biased estimates. The problem can be alleviated by using variables that control for efficiency differences of this kind. As suggested in Binswanger (1974: 381) the inclusion of an efficiency index to the right-hand side of (5) would be the first best solution but most studies actually use second best approxima-

tions like dummy variables for industries or regions. We shall the second solution in the forthcoming sections.

The third restrictive assumption concerns the exogeneity of wages. Hamermesh (1992: 456–457.) argues that by choosing firms or establishments rather than sectors or branches as the units of observation the debate on exogeneity can be taken off the agenda. However, here again, the special context of the socialist or post-socialist economy calls for caution. Under monopoly power and the existence of hidden fiscal subsidies the assumption of a 'going market wage' may not be correct implying that the model is mis-specified for the study of early stages of the transition.

2.2 Data

We use a data base providing information on within-firm relative wages and the within-firm composition of employment called the Hungarian National Labour Centre's Wage Survey (WS henceforth). The WS covered nearly full samples of firms employing more than 20 workers and approximately 10% random samples of their workers in May of each year between 1992–1999. (See *Appendix 1*). In this paper the analysis is restricted to firms reporting data on at least 30 individual workers which implies that firms with less than 300 employees are automatically excluded from the investigation.[3]

The sales revenues, material costs, depreciation (D_n) and net value of fixed assets (K_n) of the firms are known from their financial reports referring to the year under examination. Labour costs (C_{in}) were calculated by aggregating individual gross wages plus social security contributions for three groups of labour[4]:

- *unskilled*: workers with less than primary, primary or vocational manual qualifications;
- *young skilled*: workers with secondary or higher education, and work experience lower than the median;
- *older skilled*: workers with secondary or higher education and work experience equal to or higher than the median.[5]

[3] A limit of 0 individual observations per firm was also tested with similar results.

[4] To be more precise: the group-specific sum of gross wages was multiplied with 1 plus the firms's social security tax ratio calculated as the ratio of gross wages cum contributions to gross wages, taking all figures from the firm's financial report.

[5] Work experience was approximated as age minus years in school required for achieving the highest educational attainment minus 6 years. The shortcomings of the approximation are admitted. The median experience was nearly constant at 21-22 years throughout the observed period.

The firm's total costs were defined as $C_n = \sum_i C_{in} + D_n$, the sum of (monthly) labour costs plus (monthly) depreciation and the cost shares were calculated by dividing the cost bills of the three labour groups and depreciation with total costs, respectively. Thus for the three equations of (5): $s_{in} = C_{in} / C_n$, $\mathrm{i} = 1,2,3$. Factor prices were defined as expenditure per quantity employed by the firm, that is, average labour costs in the three groups of labour ($p_{in} = C_{in} / L_{in}$, L_{in} being number of workers of type i in firm n, i = 1,2,3), and depreciation per the net value of capital ($p_{4n} = D_n / K_n$). Scale effects were tested by incorporating log sales revenues net of material costs ($\ln Y$) to the right-hand side of (5). Non-neutral differences in efficiency were controlled by the use of a dummy for majority foreign ownership (F) as a proxy for highly productive technologies and potential skill-bias. Majority foreign ownership was defined as a foreign share exceeding 50% in equity. Data on ownership were available from 1992 onwards.

The most severe risk a labour demand model should face in the stage of estimation is the mismeasurement of capital stock and capital costs. The stock figure and the depreciation rate are poor measures in general but the measurement problem is further aggravated in the particular Hungarian case by at least two factors. First, the capital stock was revalued in many firms in the early years of the transition calling into question the validity of the reported figure before (or after) the revaluation. Second, Hungarian enterprises, small firms in particular, often report incredible depreciation rates as discussed in detail in Kőrösi (1998).

Unfortunately we can not directly check the validity of our data on capital. What we can try to do is estimating the model with and without capital costs. In the latter case the cost shares are to be expressed taking *labour costs* as unity and the model should be reformulated by dropping a further equation and adjusting the constraints. In case our data on capital costs have something to do with reality dropping them is expected to change the parameters for labour subgroups not separable from capital. In case the data on capital costs are completely meaningless (or capital and labour are separable) we expect no change in the estimates.

In the Interim Report we made a comparison between the two kinds of estimates. Then we concluded that we should rely on the results of the model with capital. Because of two reasons. (i) The first was the observation of close to zero cross-elasticities in cases when unskilled labour – the group with the highest elasticity of substitution with capital – was one of the labour inputs. This observation can be interpreted as an omitted variable

problem. (ii) The second signal was that by ignoring capital we receive counter-intuitive estimates for the own-wage elasticity of unskilled labour both in absolute and relative terms.

2.3 Results

The estimation results together with the own-price and cross-price elasticities calculated at the sample means of the cost shares are presented in *Appendix* 2. The movement of the elasticities between 1986 and 1999 are shown at *Figures 1–4*.

Table 1 summarizes the evolution of the cost shares over time. The footprints of the transformational recession are clearly shown by the share of depreciation in total costs falling from 57.2 to 20.8 per cent between 1986 and 1992. The rising share after the low point in 1992 hints at the rebuilding of the capital stock following the years of collapse.

Table 1

The share in total costs of unskilled, young skilled and older skilled labour and capital 1986-99

Year	UNSKILLED	Young skilled	Older skilled	Capital
	Total costs = 1			
1986	0,290	0,078	0,061	0,572
1989	0,356	0,108	0,098	0,438
1992	0,408	0,186	0,197	0,208
1993	0,381	0,190	0,199	0,230
1994	0,309	0,212	0,232	0,247
1995	0,319	0,191	0,218	0,272
1996	0,318	0,184	0,202	0,297
1997	0,298	0,182	0,189	0,331
1998	0,292	0,199	0,205	0,305
1999	0,268	0,212	0,195	0,324

Year	Unskilled	Young skilled	Older skilled	Capital
	Labour costs = 1			
1986	0,677	0,181	0,142	–
1989	0,634	0,191	0,175	–
1992	0,515	0,235	0,249	–
1993	0,494	0,247	0,258	–
1994	0,411	0,281	0,308	–
1995	0,437	0,263	0,300	–
1996	0,451	0,262	0,287	–
1997	0,445	0,272	0,283	–
1998	0,420	0,286	0,295	–
1999	0,396	0,314	0,289	–

Within labour costs the share of expenditures on unskilled workers has been almost continuously falling from 67.7 % in 1986 to 39.6% in 1999.

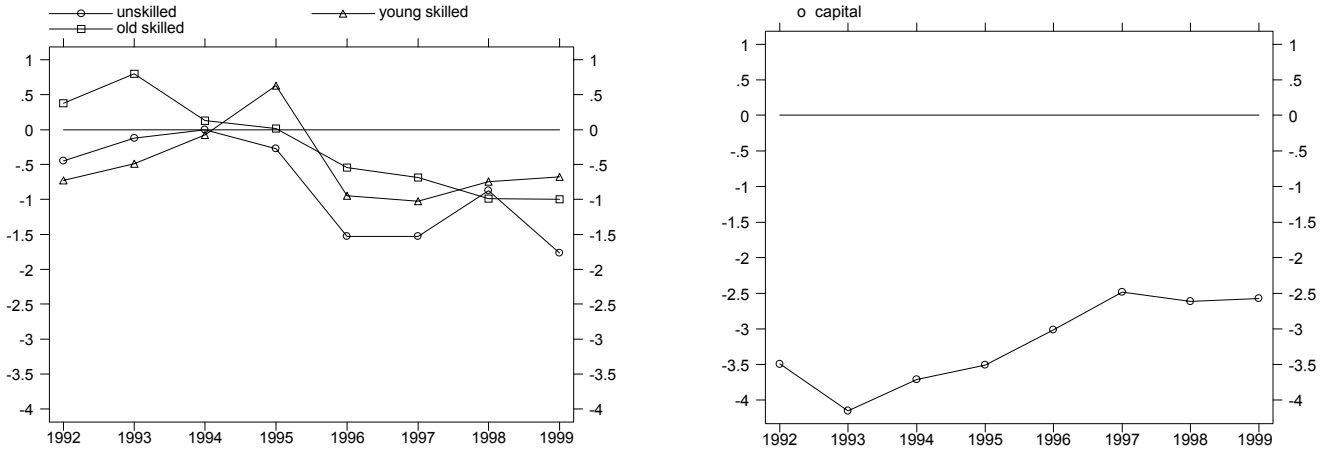
The nearly identical shares of the young and the old within the skilled subgroup follows from the definition of the splitting line.

2.3.1 *Own-price elasticities*

The estimates, perhaps surprisingly, suggest that elasticities were close to zero in 1992–95 with even positive values occurring in some groups and some years. After 1995 the elasticities were moving downwards and settled down in a range of –0.5/–1.0 for skilled and –1.0/–1.5 for unskilled workers. The own-price elasticity of capital (in absolute terms) followed a slightly decreasing path in 1994-96, and stabilizing at a level of about –2.5 in 1997–99.

Figure 1

Estimated own-price elasticities, 1992-1999



labour *capital*

At the end of the transition period the own-wage elasticity of unskilled labour was found to be the highest (in absolute terms) which is consistent with the theoretical predictions and the bulk of empirical findings from Western market economies. (See an overview of translog cost function estimates on page 460 of Hamermesh (1991)). The moderately negative own-wage elasticities at the end of the transition period is consistent with the expectations.

The shifting of the own-wage elasticities towards zero around and after the transformational recession does not come as a surprise – it seems to us as a natural consequence of rapid downsizing and fast structural change. Firms cutting their output and staff by two-digit percentages, or just starting their operations, cannot adjust factor quantities to relative factor prices instanta-

neously. In the turbulent period of the transition many of them may have been far from the optimum supposed to prevail in our cross-section model.

The time path of the own-price elasticity of capital shown at *Figure 1* is undoubtedly puzzling. For the moment we have no explanation as to why the demand for capital services was more responsive to its costs in 1992-96 than later and why did we get relatively high estimate for each year of the observed period.

2.3.2 *Cross-price elasticities*

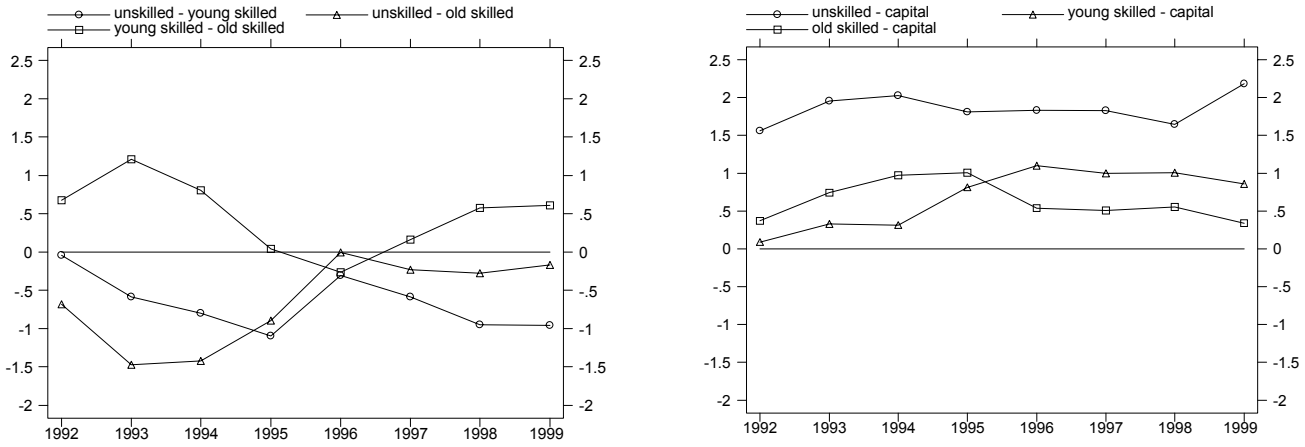
The estimated cross-price elasticities follow clear patterns. Unskilled and skilled labour are *p-complements*. Though the cross-elasticities ε_{12} and ε_{13} are both negative throughout the observed period complementarity of unskilled and older skilled labour became very weak after 1995.

Skilled-old and skilled-young workers appears to be *p-substitutes.* It should be noted that the γ_{23} parameters from which the ε_{23} elasticities had been computed were statistically *insignificant* in all years between 1992 and 1994 (*Appendix 2*) therefore the validity of the estimates for these years can be called into question. By the end of the observed period we have significant parameters but they imply positive cross-elasticities of low values. Young and old skilled labour are *weak* substitutes.

Labour and capital appear as *p-substitutes* throughout the observed period with unskilled labour having the *highest* estimated cross-elasticity. It should be mentioned again however that the parameters γ_{i4} (i =1,2,3, where j = 4 stands for capital, K) were obtained from the constraints of formulae (3b) and hence affected by the significance level of *all* the γ_{ij} $(i, j = 1,2,3,4)$ parameters. The insignificance of the γ_{23} coefficients prior to 1995 therefore has adverse implications for the reliability of the cross elasticities with capital. In the later stage of the transition when the γ_{23} values were either significant, or insignificant and small as in 1999, the values of cross elasticities with capital are more reliable. They suggest that *unskilled labour can be relatively easily substituted with capital while skilled labour is more or less separable from it*.

Figure 2

Estimated cross-price elasticities, 1992-1999



labour-labour *labour-capital*

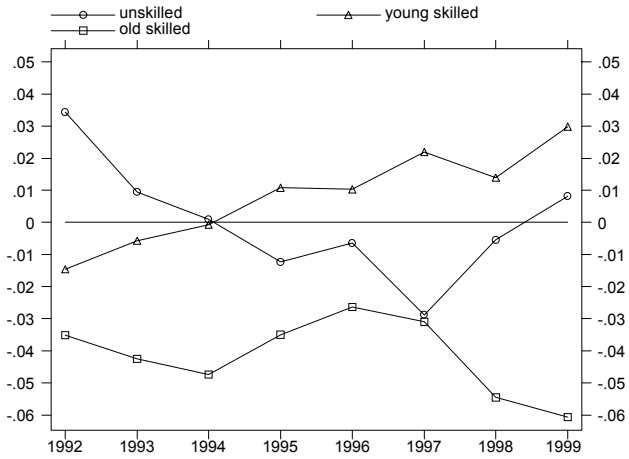
2.3.3 *Non-neutral differences in efficiency*

Figure 3 shows how the coefficients γ_{iF} $(i = 1,2,3,4)$ of the foreign ownership dummy (*F*) changed over time. Foreign-owned firms had relatively high demand for unskilled workers holding relative factor prices constant in 1992–93 but the difference between them and domestic firms changed sign in 1994–97 and virtually disappeared in 1998–99. The break in the trend is probably explained by the expansion of foreign firms in the field of industrial mass production at the end of the decade.

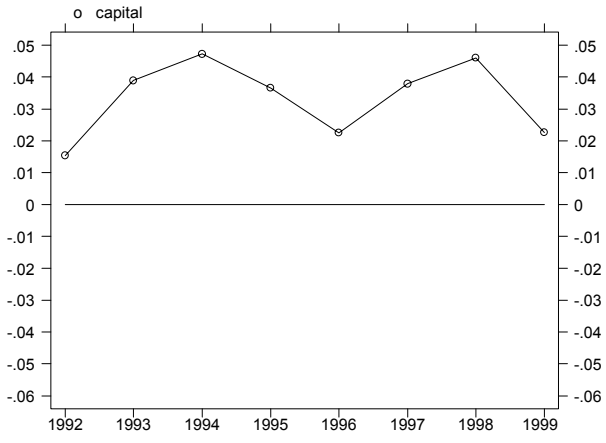
Foreign firms' demand for young skilled workers was on the rise throughout 1992–99 and was slightly higher than the average at the end of the period while the demand for skilled-old workers was relatively low and basically decreasing over time. In contrast to the case of the unskilled and the young skilled the parameters for older skilled workers were significant in all years (*Appendix 2*) suggesting that this subgroup had a share lower by 4-7 percentage points at given factor prices in foreign-owned firms. Their demand for capital was higher in the same time.

Figure 3

Parameters for foreign ownership, 1992-1999



labour



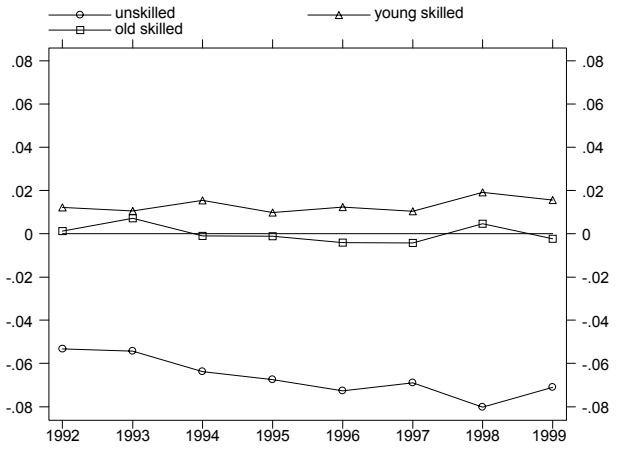
capital

2.3.4 *Scale effects*

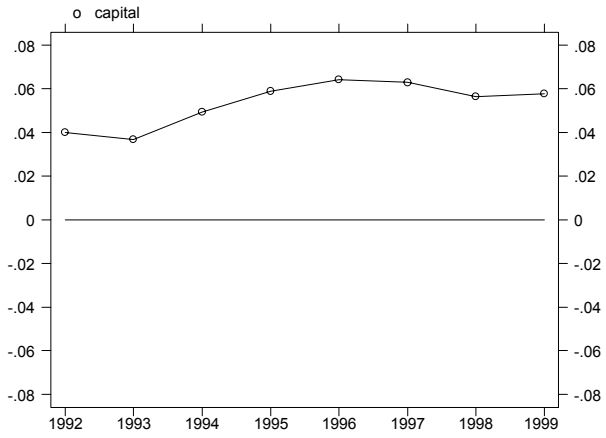
The coefficients for ln*Y* shown at *Figure 4* suggest no scale effect with respect to the two subgroups of skilled labour. Larger firms had lower demand for unskilled labour and higher for capital throughout the observed period with no significant fluctuations in the estimated parameters.

Figure 4

Scale effects (parameters for ln*Y*), 1992-1999



labour



capital

3. USING LABOUR DEMAND ESTIMATES IN THE EVALUATION OF THE POTENTIAL CONSEQUENCES OF THE RISE IN MINIMUM WAGE

Minimum wage was doubled in nominal terms in two years between 1999 and 2002 in Hungary. The 1999 minimum wage – 27,8 percent of the mean earnings of that year – was raised January 2000 to 25.000 forints, than a huge rise followed in January 2001: the new minimum wage was set at a level of 40.000 forints (41,5 per cent of the average earnings of that month). In January 2002 a further big rise took place: the new minimum wage was set at 50.000 forints, which corresponds to 45,5 percent of the actual mean earnings. It is hard to believe that an intervention of such a scale has no serious employment consequences.[6] However, governmental agencies asserted exactly the opposite by referring to some "direct" evidence: first of all to the low percentage of those recent job losers who attributed the loss of their jobs to the rise of minimum wage when interviewed. The aim of this section is to evaluate the impact of the minimum wage hike on labour demand. While economic theory sheds some light on the effects of raising the minimum wage the actual orders of magnitudes are far from being clear. In what follows we try to put together two pieces of information: (i) the wage elasticities estimated from the multi-factor demand system of the former section of this paper, and (ii) the actual earnings distribution(s) of the year 1999 when this "social experiment" took place. As more recent (individual level) wage data are not available the changing shape of earnings distribution – what we predict[7] – has to be confirmed in the near future by the Wage Survey data of the years 2001 and 2002. Using the estimated (own-wage) elasticities for the type i of labour (ε_{ii}) predicted drops in demand will be calculated by the formula:

$$\frac{\Delta\hat{L}_i}{\hat{L}_i} = \varepsilon_{ii} \frac{E(w_i \mid w_i = w'_{\min} \text{ if } w_i < w'_{\min}) - E(w_i \mid w_i = w_{\min} \text{ if } w_i < w_{\min})}{E(w_i \mid w_i = w_{\min} \text{ if } w_i < w_{\min})}, \quad w'_{\min} > w_{\min} \quad (7)$$

Minimum wages are taken at real value as if they were raised to the same percentages of the 1999 average wage as they are related (at current value)

---

[6] Minimum wage law covers all workers in Hungary who have employment contracts. *No sector is uncovered.*

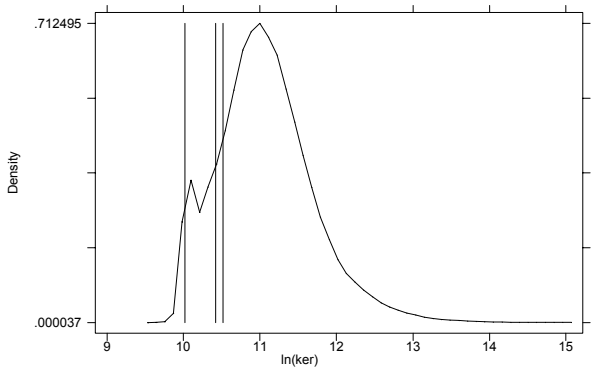
[7] An evidence for this is presented in *Appendix 3*. Wage distributions stem from two subsequent surveys (from 1994 and 2001). Samples covered workers who entered employment from the UI register within the span of a month in 1994 and 2001. Wages are subjective estimates of the entrants as their own maximum and minimum expected wages. Wages underlying the distribution were calculated as averages of these expected minima and maxima.

to the mean wage of that year and month (January of 2001 and 2002) when they were really set.[8]

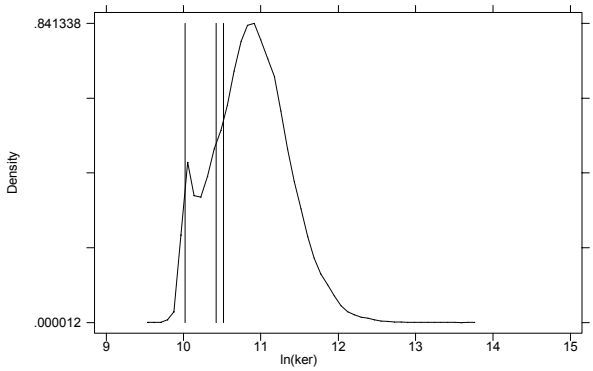
The higher the own-wage elasticity of a certain type of labour and the more people are affected by the differential of the old and the new minimum wage standard the stronger are the expected demand responses. The population most in danger are low educated people, – for both reasons: their demand is most responsive to wages (*Figure 1*), and most of those workers whose wages have to be increased by the new minimum wage standard are unskilled-low wage people. *Figure 5* clearly proves this.

Figure 5

The distribution of log wages in the non-budget sector by educational attainment in 1999 (the distributions are split by the subsequent minimum wage levels; minimum wages are calculated at 1999 real value)

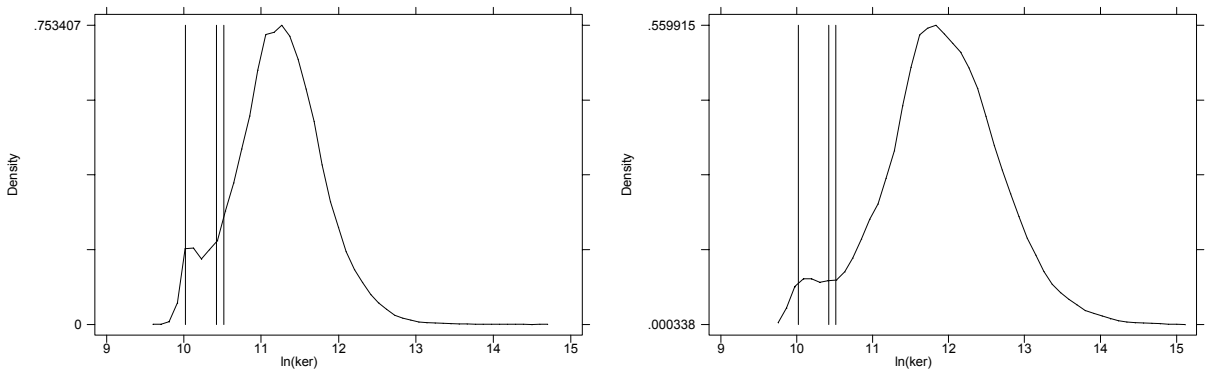


Non budget sector as a whole



Unskilled workers: completed/incomplete primary school or vocational training school

[8] The minimum wage of the year 2001 (forints 40.000; 41,5 % of the mean wage of January 2001) is 33.607 forints at 1999 value, the minimum wage of the year 2002 (forints 50.000; 45,5 % of the mean wage of January 2002) is 36.847 forints at 1999 value.



Secondary school College

Using the formula (7) and the estimated own-wage elasticities calculated from the multi-factor demand model of section 1, the predicted decrease in demand is given by *Table 2*.

Table 2

Predicted decrease in labour demand in the non-budget sector, 1999[a]

Type of labour	% in total labour force in 1999[a]	Own-wage elasticity, in 1999[b]	minimum wage: HUF 22.500 => HUF 40.000[c]		minimum wage: HUF 22.500 => HUF 50.000[c]	
			$\Delta w / w$ (%)	$\Delta L / L$ (%)	$\Delta w / w$ (%)	$\Delta L / L$ (%)
Unskilled	60,6	–1,768	2,18	–3,85	3,38	–5,98
Young skilled	19,7	–0,647	0,61	–0,39	0,91	–0,59
Older skilled	19,7	–0,997	0,37	–0,37	0,48	–0,48
Total	100,0	–	1,51	–2,48	2,32	–3,83

[a] Calculated from the individual data file of the Wage Survey, 1999.

[b] See *Figure 1* and *Appendix 2*

[c] The minimum wage of 2001 (Ft 40.000) is taken at 1999 value as Ft 33.607, The minimum wage of 2002 (Ft 50.000) is taken at 1999 value as Ft 36.847.

The large sample size of the Wage Survey (we have more than 100.000 workers in the non-budget sector) allows us to repeat this exercise by different regional units. As low educated people are distributed very unevenly in the country we can expect very diverse demand consequences of such a drastic measure as doubling the minimum wage in a short, two year, period.

We broke down the sample of the Wage Survey by two regional dimensions: by counties (20 units included Budapest) and types of settlement (county capital, other towns, and villages), and distinguished 57 regional units. As educational and wage differentials appeared to be extremely large no wonder that the drops in demand implied by the rise in minimum wage dispersed in a very broad range. In *Appendix 4* several wage distributions are shown. While some regions are affected modestly, others – mostly villages in the Northeastern part of the country and in the Hungarian Plain – are hit extremely severely by the potential employment consequences of the rise of minimum wage. In some cases the doubling of the minimum wage simply splits the distribution in two equal parts at the mode. As the employment effects vary mostly with the share of low educated (unskilled) workers, we calculate predicted changes in labour demand for two categories of low educated workers: for those whose educational attainment is not higher than completed primary school, and for those who completed vocational training school. Results are presented by *Figure 6*.

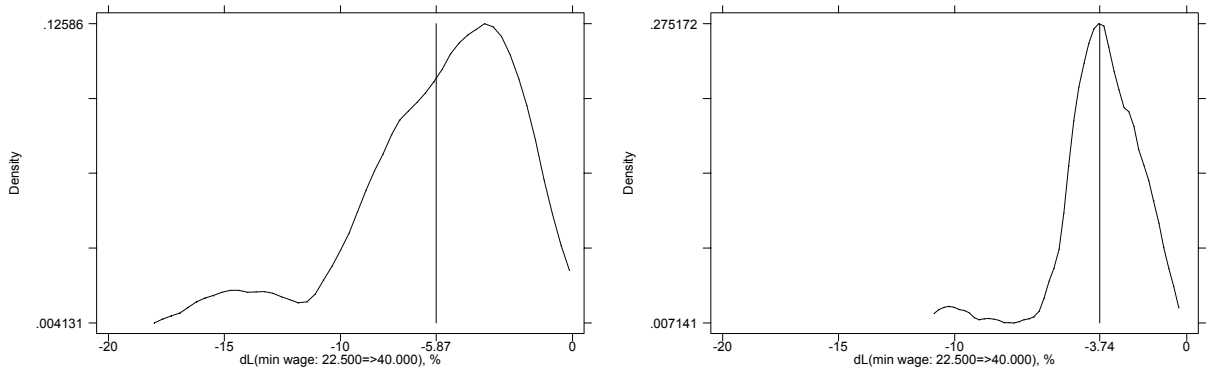
As the upper two panels of *Figure 6* show demand for low educated labour is expected to decrease by the rise of minimum wage very unevenly in the country. In some regions more than a 10 per cent drop is expected for workers with completed or incomplete primary school and more than 5 percent for those who completed vocational training school. The rise of the minimum wage does increase regional employment inequalities. No matter that the further rise in minimum wage in 2002 represents a 10 per cent hike relative to the former level – 50.000 forints relative to 40.000 forints in nominal terms; and 36.000 forints relative to 33.000 forints in real terms – in some regions this affects so many low wage workers that it may aggravate the employment crises on a massive scale. As the two lower panels of *Figure 6* prove, this can add another 3–5 per cent employment decline to the existing large (10-15 percent) decline.

The *actual* degree of employment decrease may be less than the *predicted*. Several forces can mitigate the employment decline: (i) Employers may not comply with the new minimum wage standard. Formally they comply with the rules but informally they can evade it. As workers whose jobs are in danger are partners in the evasion of the minimum wage standards, in many cases they can do this. Employers can make part-time contracts and actually require full time services; they can set performance pay standards

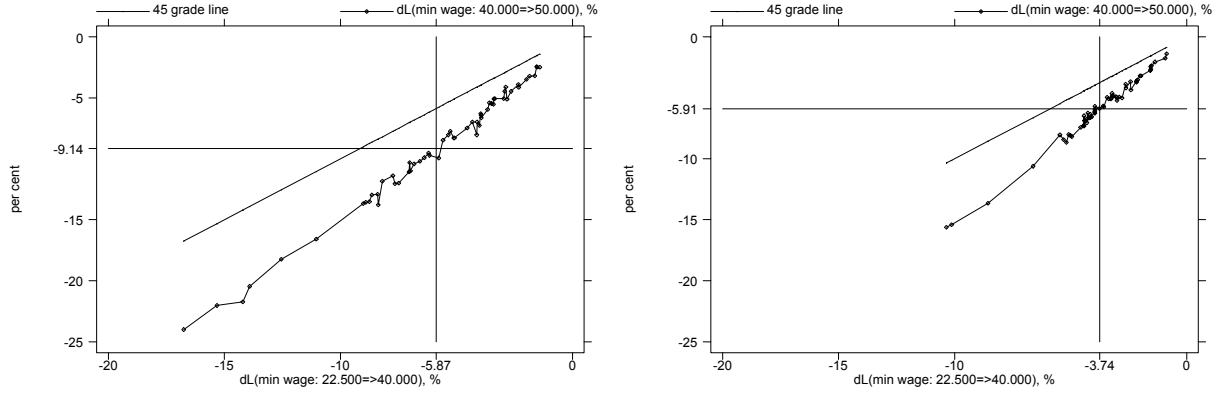
Figure 6

Predicted decrease in labour demand implied by the rise of minimum wage by educational attainment and regional types (N=57)

Primary school | Vocational training school



the distribution of regional types by $\ddot{A}D_L$ implied by the rise of minimum wage: 22.500 => 40.000



the increase of $\ddot{A}D_L$ by the further rise in minimum wage (40.000 => 50.000) by regional types

with never can be met at 100 per cent, etc. Actual hours and actual performance are hard to control from outside. (ii) If the new minimum wage is effective, it reduces the within-firm wage dispersion that may weaken the incentive effects of wages (particularly if some *relative* performance measure is applied). In these cases there is a pressure of raising the wages in the range above the minimum (and mean) wage. Thus relative wages can get closer to the initial relativities. (iii) As minimum wage increase changes the wage/unemployment benefit ratio it affects labour supply as

well. Unemployment consequences can be lessened by the shift in labour supply.

We have no information on the strength of these effects. However the predicted demand effects are so strong, that we expect *actual* employment decline on aggregate as well as an increase of regional employment inequalities due to the minimum wage hike of 2001. We have two pieces of evidence.

(i) Recent *employment statistics* (based on quarterly labour force surveys) report the changes in employment by the combined criteria of sector (agriculture, industry, and services) and six larger regional units of the country. Based on these data, percentage changes in employment $((L_t - L_{t-1}) / L_{t-1})$ [9] are regressed on percentage changes in wages $((w_t - w_{t-1}) / w_{t-1})$ [10] implied by the minimum wage hike. If our predictions are right we expect statistically significant and negative effect of *dw* on dL. This will be tested first.

(ii) A second piece of evidence comes from the community level *unemployment statistics* based on the registers of the local labour centres (which are responsible for registering unemployed people and paying unemployment benefits to them). Unemployment rates[11] are calculated for 150 micro-regions and the time paths are followed by the changes of the quarterly rates throughout the 1991–2001 period (44 quarters). Expressing mean unemployment rates of the different (10^{th}, 9^{th}, 8^{th}, …, 2^{nd}, 1^{st}) deciles in percentages of the median at each period of time, we get a picture of the time path of the relative unemployment rate differentials. If the time paths of the deciles (particularly of the upper deciles) *break* at some point of time after the end of 2000 (just before minimum wage was raised on 40.000 forints) and unemployment rate differentials start to increase *relative to the time trend* we can suspect the impact of minimum wage hike. This will be our second test.

Look at first test 1. Percentage changes in employment are regressed on the percentage changes in wages implied by the minimum wage hike. As *Figure 7* indicates, we have statistically significant and quite strong negative effect: a one percent wage increase implied by the rise in minimum wage has 8,5 per cent negative impact on the employment level. In the relevant

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[9] See: Labour Force Survey, Quarterly Bulletin, Quarter 4 of 2001, Central Statistical Office, Budapest, p. 36; t = October-December of 2001, $t - 1$ = October-December of 2000.

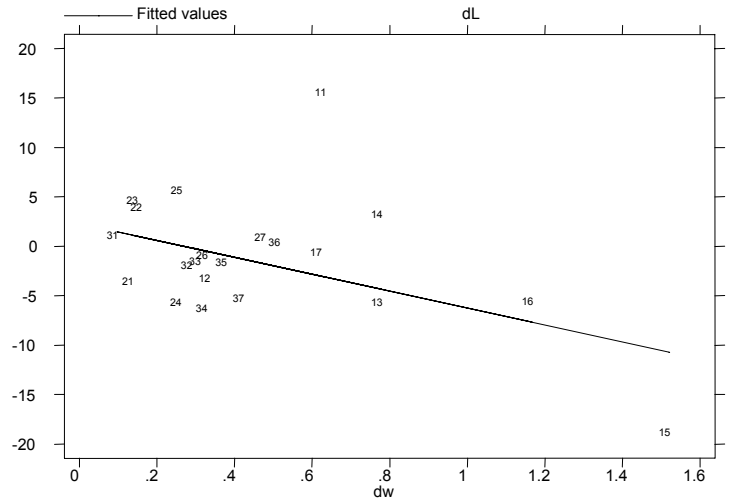
[10] *Source*: Wage Survey, 1999. *dw*/*w* is calculated as it is indicated in the right hand side of formula (7).

[11] Unemployment rates are calculated by the ILO standards: $u = U/(E + U)$, E = # of employed, U = # of un-employed.

range (about ½ percent rise in wages) we find about 4 per cent employment decline.

Figure 7

Employment effects of the rise in mean wages implied by the rise of minimum wage between the last quarters of 2000 and 2001, %
($N = 21$ regional x sectoral units)

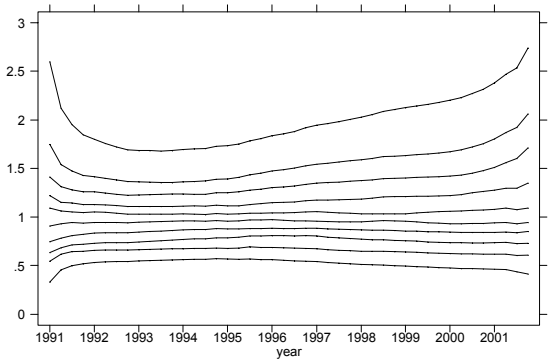


$$dL = 2.31 - 8.56\ dw,\ N = 21\ \text{adj}R^2 = 0.175$$
$$(1.06)\quad (-2.29)$$

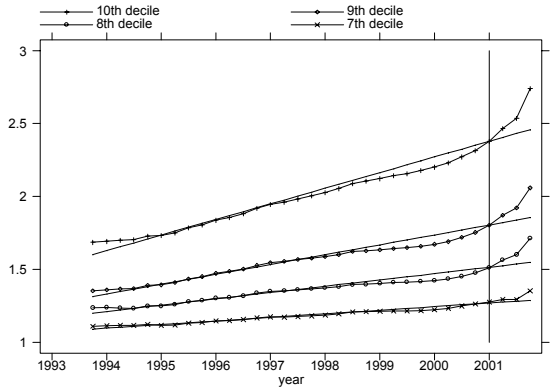
The second test serves to check the proposition that it is not completely unjustified to blame the 2001 minimum wage hike for the increase of regional unemployment rate differentials. *Figure 8* presents mean quarterly rates of 150 districts of the country by distinguishing 10 deciles by local unemployment rates. (As indicated before quarterly rates are divided by the given quarter's median unemployment rates.) The upper left panel of the Figure shows the story of the regional unemployment rate differentials over the whole 11-year period. What is important from the present paper's aspect is that after a two-years long period of stable regional differentials (1992–93) the dispersion was starting to grow at an almost steady rate from the end of 1993 onwards. To test this linear time trends were fitted on each decile from the last quarter of 1993. As the two lower panels indicate, linear trends make quite a good job in predicting the time pattern of certain deciles. In these two graphs we drew standardised residuals – (actual – trend)/trend – in order check this. Standardised residuals are fluctuating within a (–5,+5) per cent bound for the first seven deciles.

Figure 8

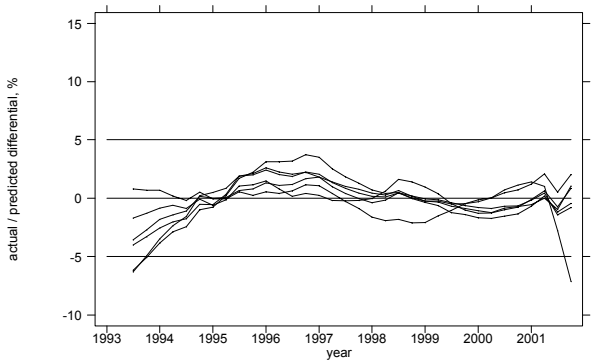
Relative unemployment rate differentials of micro-regions
(N = 150), 1991-2001 (by deciles)



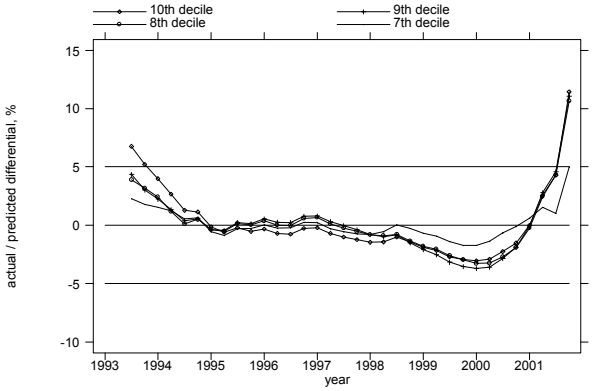
(a) Relative rates = unemployment rates divided by the rate of the median district



Relative rates (see: panel (a)) and linear trends for the 7^{th}, 8^{th}, 9^{th}, 10^{th} deciles



Standardised residuals = (actual-trend)/trend for the 1^{st}, 2^{nd},…, 6^{th} deciles



Standardised residuals = (actual-trend)/trend for the 7^{th} , 8^{th}, 9^{th}, 10^{th} deciles

However, he story of the upper three deciles is completely different. The time trends predict the actual paths quite well until the end of 2000. From the first quarter of 2001 onwards we have a dramatic upward deviation from the 7 year-long trend in the upper three deciles. It is hard to avoid a conclusion that this sudden deviation from the former path (particularly in those regions where we predict exactly the same development) has something to do with the shock of minimum wage hike of January 2001.

A final comment. The exercises made in this section cannot stand for a

proper analysis of the actual employment impacts of rises in minimum wage. Future research is required to discover the full scale of impacts of this dramatic "non-natural experiment". Individual wage and employment data sets of the years 2000, 2001, 2002 are required to do this job. The present writers plan to do this job in the near future. However the present estimates seem to us so robust that we cannot avoid the conclusion: the minimum wage hike of 2001 and 2002 contributed to the employment decline in Hungary from the end of 2001 in a non-trivial extent and increased regional inequalities. May be this price is not too high relative to the benefits. But this dilemma should be at least an issue on the political agenda.

4. CONCLUSIONS

This paper analyses changes in the demand for unskilled, young skilled, and older skilled workers during the post-communist transition in Hungary. Systems of cost share equations derived from the translog cost function are estimated for cross-sections of large firms observed in the period 1992–99. Following the 'transformational recession' the own-price elasticities of labour and capital were stabilized at levels observed in several developed market economies. Unskilled and skilled labour are estimated to be *p*-complements, and younger and older skilled workers *p*-substitutes. Capital and labour appear to be *p*-substitutes with unskilled labour having the highest elasticity of substitution. Further results hint at the existence of non-negligible scale effects and the non-neutrality of technical change. Foreign firms have relatively low demand for older skilled workers.

The high own-price elasticity of unskilled labour and its high degree of substitution with capital warns that the unskilled labour market remains a fragile segment of the economy. In recent years the increasing budget revenues, the decreasing rate of unemployment and the wish to catching up to EU wage levels encouraged the Hungarian government to change its preferences in employment policies. Unemployment benefits were curtailed, job creation subsidies were extended and a decision was made to double the minimum wage in the course of only two years. The first step was made in 2001 by an increase of 60%, which was followed in 2002 by another big increase. In view of the results presented in this paper these policies should be considered mistaken unless they rest on the assumption that the supply side effects of lower benefits and higher minimum wages will offset the adverse impact on the demand side. A policy decreasing the user cost of capital and increasing the relative wage of the least quali-

fied workers is expected to diminish the demand for unskilled labour. If our estimates are correct wage subsidies for the unskilled can be advised instead for politicians who care about the 'lower segments' of the society.

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APPENDIX 1

DRAWING FIRM SAMPLES FROM THE WAGE SURVEY

The National Labour Centre's Wage Survey (WS) was carried out in May 1986, May 1989 and each May since 1992. It contains data of about 150,000 workers employed in 6,000 to 12,000 firms, depending on year. The sampling procedure is two-step. At the first step branches (geographically distinct plants, shops or offices of enterprises) are selected. At the second a 10% random sample of full-time employees is drawn within each branch.

In principle the WS covers all business firms employing more than 10 workers (20 workers prior to 1995), and all institutions of the budget sector excluding armed forces. In practice smaller firms are under-represented due to non-response. Business firms should select an approximately 10% random sample of their full-time employees. (The selection is based on the workers' day of birth). In small firms the sampling quota is higher. Budget institutions report data for all workers. Individual weights (w_1) are included in the original data files. The sum of individual weights across a firm is equal to the number of the firm's employees. The samples for this research were drawn in the following way:

(i) Firms reporting less than 30 individual observations were dropped.

(ii) The sum of weights (employment), the weighted sum of gross wages (wage bill), and average wages were calculated for each of the three skill groups.

(iii) Gross average wages and wage bills were multiplied by the firm's social security tax ratio.

(iv) Depreciation, output (sales revenues net of material costs and depreciation), and data on ownership were drawn from the firm's financual report.

(v) The cost shares and factor prices were calculated as described in the text.

Since relatively small „large firms" (those employing 300–400 workers) have a relatively high risk of having less than 30 workers born on certain days of the week they are at risk of dropping out from the sample. To correct for this sort of random attrition the ratio of firms in the restricted sample to firms in the WS was attached to firms in the following size categories: 300–1000, 1000–3000, 3000–.

The calculations presented in this draft exclude firms with negative value added, those with zero share of unskilled workers, and those with a depreciation rate exceeding 1. The total number of firms excluded for these rea-

sons is summarised in the table below. The bias resulting from exclusions will be discussed in later versions.

	$s_1 = 0$	d>1	Y < 0
1992	0	2	20
1993	1	0	20
1994	0	1	20
1995	2	0	0
1996	2	3	11
1997	7	0	10
1998	6	0	14
1999	10	3	12

APPENDIX 2

ESTIMATION OF MODEL (5) – (6)

(A) Test statistics of the equations

Equation	Nobs	RMSE	„R-sq"	Chi2	Dependent var. Mean	Dependent var. St. dev.
			1992			
Unskilled	571	,16934	0,2099	149,7442	,408	,194
Young skilled	571	,09648	0,1336	100,7881	,186	,107
Old skilled	571	,10573	0,2206	187,9917	,197	,107
Capital					,208	,169
			1993			
Unskilled	541	,15673	0,2898	245,352	,380	,193
Young skilled	541	,09289	0,1432	89,19383	,190	,108
Old skilled	541	,10372	0,2278	176,3672	,198	,110
Capital					,230	,176
			1994			
Unskilled	489	,14481	0,3355	266,8304	,300	,184
Young skilled	489	,08822	0,2480	162,889	,211	,104
Old skilled	489	,10542	0,2234	157,2303	,231	,111
Capital	**				,247	,188
			1995			
Unskilled	460	,14328	0,3634	271,0769	,318	,180
Young skilled	460	,08695	0,2266	146,5132	,191	,094
Old skilled	460	,10292	0,2115	156,5314	,218	,098
Capital	**				,270	,190
			1996			
Unskilled	458	,15812	0,2947	196,5087	,317	,196
Young skilled	458	,09325	0,1290	102,7404	,184	,101
Old skilled	458	,09755	0,2010	138,7752	,201	,100
Capital	**				,296	,185
			1997			
Unskilled	605	,14637	0,3577	353,6924	,297	,185
Young skilled	605	,08329	0,1530	139,5373	,182	,090
Old skilled	605	,08625	0,1720	140,4203	,189	,087
Capital	**				,330	,195

Equation	Nobs	RMSE	„R-sq”	Chi2	Dependent var.	
					Mean	St. dev.
			1998			
UNSKILLED	455	,15501	0,3753	281,8355	,291	,202
Young skilled	455	,08859	0,2343	151,7401	,198	,115
Old skilled	455	,09740	0,1843	115,1044	,204	,100
Capital	**				,304	,190
			1999			
Unskilled	436	,14818	0,4371	340,7512	,267	,200
Young skilled	436	,09362	0,2660	174,1988	,212	,128
OLD SKILLED	436	,09854	0,1911	103,9811	,195	,100
Capital	**				,324	,201

Number of estimated parameters: 4 in 1986–89, 5 in 1992–99.
Number of constraints: 3
All Chi2 tests significant at 0.0000

(B) Parameter estimates

	Parameter	St. error	Z	Sign.
		1992		
Unskilled (1)				
ln(p1)	,1667569	,0335702	4,967	0,000
ln(p2)	-,0792427	,0171395	-4,623	0,000
ln(p3)	-,1353938	,0170666	-7,933	0,000
Foreign	,0343055	,0272017	1,261	0,207
ln(Y)	-,0532346	,0070322	-7,570	0,000
_cons	1,386346	,1868823	7,418	0,000
Young skilled (2)				
ln(p1)	-,0792427	,0171395	-4,623	0,000
ln(p2)	,1264968	,0141826	8,919	0,000
ln(p3)	-,0118399	,0098114	-1,207	0,228
Foreign	-,0146061	,015721	-0,929	0,353
ln(Y)	,0121137	,0040586	2,985	0,003
_cons	-,2938378	,1168805	-2,514	0,012
Old skilled (3)				
ln(p1)	-.1353938	.0170666	-7.933	0.000
ln(p2)	-.0118399	.0098114	-1.207	0.228
ln(p3)	.1729956	.0130838	13.222	0.000
Foreign	-.0350703	.0171485	-2.045	0.041
ln(Y)	.0010743	.0043996	0.244	0.807
_cons	-.1595204	.1107491	-1.440	0.150
		1993		
Unskilled (1)				
ln(p1)	,2183059	,0362139	6,028	0,000
ln(p2)	-,1149678	,0191758	-5,995	0,000
ln(p3)	-,1870832	,0188024	-9,950	0,000
Foreign	,0094524	,0197926	0,478	0,633
ln(Y)	-,0543843	,0070678	-7,695	0,000
_cons	1,809777	,2138859	8,461	0,000
Young skilled (2)				
ln(p1)	-,1149678	,0191758	-5,995	0,000
ln(p2)	,1363761	,0162796	8,377	0,000
ln(p3)	,0079007	,011712	0,675	0,500
Foreign	-,0057658	,011814	-0,488	0,626
ln(Y)	,0106247	,0042425	2,504	0,012
_cons	-.2436869	.1330338	-1.832	0.067

Old skilled (3)				
ln(p1)	-,1870832	,0188024	-9,950	0,000
ln(p2)	,0079007	,011712	0,675	0,500
ln(p3)	,1908595	,0163516	11,672	0,000
Foreign	-,0425717	,0131188	-3,245	0,001
ln(Y)	,0070301	,0046881	1,500	0,134
_cons	-,0630462	,1394038	-0,452	0,651
	1994			
Unskilled (1)				
ln(p1)	,2132563	,0339015	6,290	0,000
ln(p2)	-,1178909	,0168463	-6,998	0,000
ln(p3)	-,1736723	,018304	-9,488	0,000
Foreign	,0009093	,0182654	0,050	0,960
ln(Y)	-,0637308	,0071532	-8,909	0,000
_cons	1,795071	,214238	8,379	0,000
Young skilled (2)				
ln(p1)	-,1178909	,0168463	-6,998	0,000
ln(p2)	,1634983	,0140875	11,606	0,000
ln(p3)	-,0097392	,0108352	-0,899	0,369
Foreign	-,0007203	,0112099	-0,064	0,949
ln(Y)	,0154116	,0043045	3,580	0,000
_cons	-,3373976	,1273682	-2,649	0,008
Old skilled (3)				
ln(p1)	-,1736723	,018304	-9,488	0,000
ln(p2)	-,0097392	,0108352	-0,899	0,369
ln(p3)	,1851029	,0163851	11,297	0,000
Foreign	-,0474108	,0133306	-3,557	0,000
ln(Y)	-,001014	,0050672	-0,200	0,841
_cons	,1347838	,1403291	0,960	0,337
	1995			
Unskilled (1)				
ln(p1)	,1892055	,0355837	5,317	0,000
ln(p2)	-,1277545	,0183589	-6,959	0,000
ln(p3)	-,1319742	,0178469	-7,395	0,000
Foreign	-,01231	,0158931	-0,775	0,439
ln(Y)	-,0674899	,0076217	-8,855	0,000
_cons	1,74335	,2398341	7,269	0,000

Young skilled (2)				
ln(p1)	-,1277545	,0183589	-6,959	0,000
ln(p2)	,1776744	,0157972	11,247	0,000
ln(p3)	-,0401501	,0111128	-3,613	0,000
Foreign	,0107609	,0097379	1,105	0,269
ln(Y)	,009861	,00462	2,134	0,033
_cons	-,0315949	,1429938	-0,221	0,825
Old skilled (3)				
ln(p1)	-,1319742	,0178469	-7,395	0,000
ln(p2)	-,0401501	,0111128	-3,613	0,000
ln(p3)	,1715333	,0157397	10,898	0,000
Foreign	-,0349688	,0114298	-3,059	0,002
ln(Y)	-,0012794	,005355	-0,239	0,811
_cons	,1866839	,1476522	1,264	0,206
	1996			
Unskilled (1)				
ln(p1)	,062662	,0365162	1,716	0,086
ln(p2)	-,0764933	,0181458	-4,215	0,000
ln(p3)	-,0645314	,0174054	-3,708	0,000
Foreign	-,0064513	,0178499	-0,361	0,718
ln(Y)	-,0726571	,0081829	-8,879	0,000
_cons	1,88402	,2261915	8,329	0,000
Young skilled (2)				
ln(p1)	-,0764933	,0181458	-4,215	0,000
ln(p2)	,1180396	,0145604	8,107	0,000
ln(p3)	-,0470167	,0101782	-4,619	0,000
Foreign	,010373	,0105677	0,982	0,326
ln(Y)	,0124353	,0048098	2,585	0,010
_cons	,1201144	,1288522	0,932	0,351
Old skilled (3)				
ln(p1)	-,0645314	,0174054	-3,708	0,000
ln(p2)	-,0470167	,0101782	-4,619	0,000
ln(p3)	,1390068	,0135579	10,253	0,000
Foreign	-,0263858	,0110546	-2,387	0,017
ln(Y)	-,0040508	,0049699	-0,815	0,415
_cons	-,1188122	,1259315	-0,943	0,345

	1997			
Unskilled (1)				
ln(p1)	,0736403	,0302882	2,431	0,015
ln(p2)	-,0858464	,0144312	-5,949	0,000
ln(p3)	-,0693402	,0141624	-4,896	0,000
Foreign	-,0289112	,0149093	-1,939	0,052
ln(Y)	-,0690232	,0068762	-10,038	0,000
_cons	1,913219	,2143816	8,924	0,000
Young skilled (2)				
ln(p1)	-,0858464	,0144312	-5,949	0,000
ln(p2)	,1149044	,0120085	9,569	0,000
ln(p3)	-,0288151	,008571	-3,362	0,001
Foreign	,0218951	,008523	2,569	0,010
ln(Y)	,010367	,0038777	2,673	0,008
_cons	,0478312	,1190327	0,402	0,688
Old skilled (3)				
ln(p1)	-,0693402	,0141624	-4,896	0,000
ln(p2)	-,0288151	,008571	-3,362	0,001
ln(p3)	,1289175	,0115962	11,117	0,000
Foreign	-,0309393	,0088225	-3,507	0,000
ln(Y)	-,0042934	,0039615	-1,084	0,278
_cons	-,1675047	,1205093	-1,390	0,165
	1998			
Unskilled (1)				
ln(p1)	,1321388	,0326433	4,048	0,000
ln(p2)	-,1130624	,0155539	-7,269	0,000
ln(p3)	-,0764637	,0159774	-4,786	0,000
Foreign	-,0054252	,0168973	-0,321	0,748
ln(Y)	-,0801749	,0078952	-10,155	0,000
_cons	1,764961	,2304695	7,658	0,000
Young skilled (2)				
ln(p1)	-,1130624	,0155539	-7,269	0,000
ln(p2)	,1297768	,0128367	10,110	0,000
ln(p3)	-,017304	,0094763	-1,826	0,068
Foreign	,0139076	,0098672	1,409	0,159
ln(Y)	,0190708	,004489	4,248	0,000
_cons	-,0135275	,1300916	-0,104	0,917

Old skilled (3)				
ln(p1)	-,0764637	,0159774	-4,786	0,000
ln(p2)	-,017304	,0094763	-1,826	0,068
ln(p3)	,1215368	,0132581	9,167	0,000
Foreign	-,0544742	,0106561	-5,112	0,000
ln(Y)	,0045931	,0048532	0,946	0,344
_cons	-,1883703	,13403	-1,405	0,160
	1999			
Unskilled (1)				
ln(p1)	,0693846	,0343169	2,022	0,043
ln(p2)	-,1112451	,0165407	-6,726	0,000
ln(p3)	-,0611927	,0171847	-3,561	0,000
Foreign	,0081811	,0163805	0,499	0,617
ln(Y)	-,0709971	,0078919	-8,996	0,000
_cons	2,199005	,2472808	8,893	0,000
Young skilled (2)				
ln(p1)	-,1112451	,0165407	-6,726	0,000
ln(p2)	,1368696	,0140884	9,715	0,000
ln(p3)	-,0161106	,0104254	-1,545	0,122
FOREIGN	,0297925	,010397	2,865	0,004
ln(Y)	,0156725	,0050029	3,133	0,002
_cons	-,1113498	,1408705	-0,790	0,429
Old skilled (3)				
ln(p1)	-,0611927	,0171847	-3,561	0,000
ln(p2)	-,0161106	,0104254	-1,545	0,122
ln(p3)	,1191438	,0148532	8,021	0,000
Foreign	-,0606087	,0109334	-5,543	0,000
ln(Y)	-,0023849	,0052088	-0,458	0,647
_cons	-,2951256	,1462036	-2,019	0,044

(C) Elasticity Estimates

(i) Own price elasticities

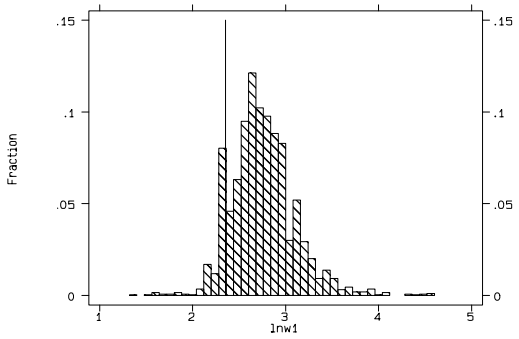
	Unskilled	Young	Old	Capital
	ε_{11}	ε_{22}	ε_{33}	ε_{KK}
1992	-0,449	-0,724	0,378	-3,493
1993	-0,120	-0,489	0,800	-4,148
1994	-0,004	-0,074	0,131	-3,712
1995	-0,275	0,630	0,018	-3,507
1996	-1,528	-0,949	-0,543	-3,014
1997	-1,528	-1,026	-0,687	-2,485
1998	-0,875	-0,746	-0,986	-2,610
1999	-1,768	-0,674	-0,997	-2,573

(ii) cross-price elasticities

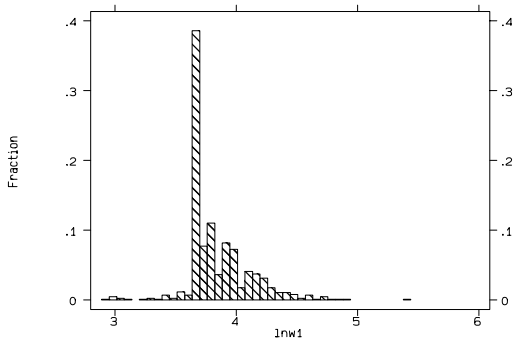
	Unskilled-young	Unskilled-old	Young-old
	σ_{12}	σ_{13}	σ_{23}
1992	-0,042	-0,683	0,678
1993	-0,587	-1,473	1,209
1994	-0,801	-1,421	0,801
1995	-1,098	-0,897	0,038
1996	-0,309	-0,007	-0,265
1997	-0,584	-0,229	0,165
1998	-0,949	-0,278	0,575
1999	-0,956	-0,170	0,612
	Unskilled-capital	Young-capital	Old-capital
	σ_{1K}	σ_{2K}	σ_{3K}
1992	1,563	0,088	0,373
1993	1,955	0,331	0,745
1994	2,024	0,314	0,970
1995	1,814	0,812	1,010
1996	1,832	1,100	0,541
1997	1,828	0,996	0,509
1998	1,646	1,010	0,555
1999	2,187	0,862	0,340

APPENDIX 3

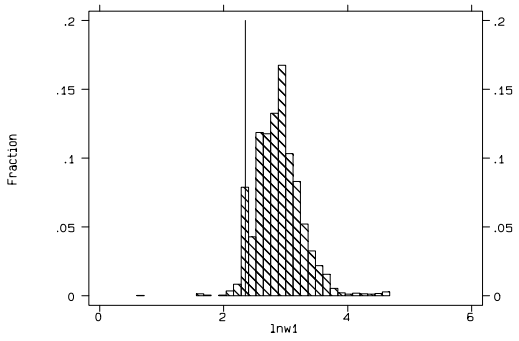
THE DISTRIBUTION OF LOG EARNINGS AFTER THE UNEMPLOYMENT, 1994, 2001[12]



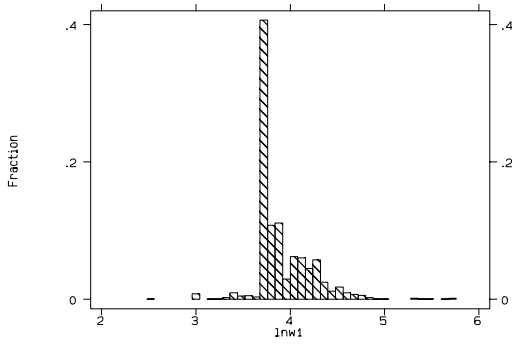
Completed primary of less, 1994



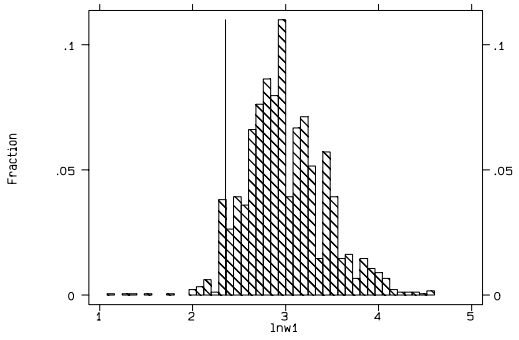
Completed primary of less, 2001



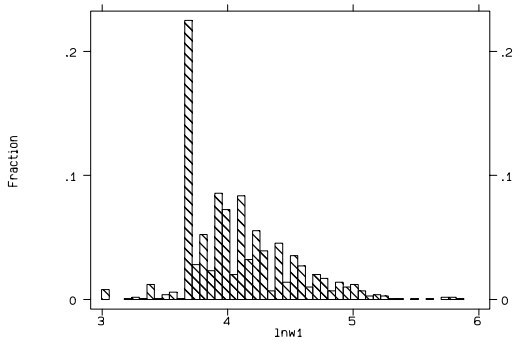
Vocational training school, 1994



Vocational training school, 2001



Secondary or college, 1994



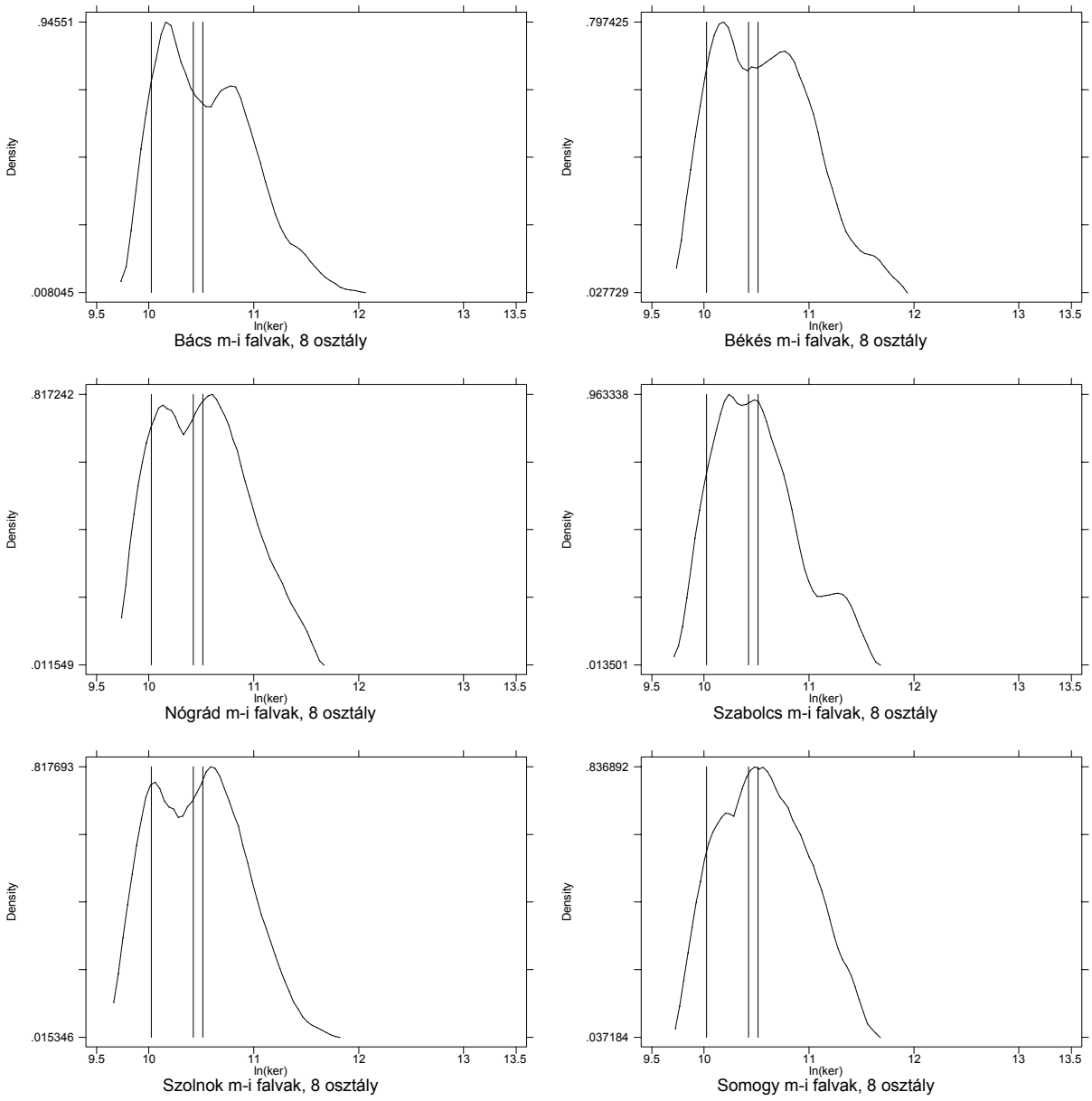
Secondary or college, 2001

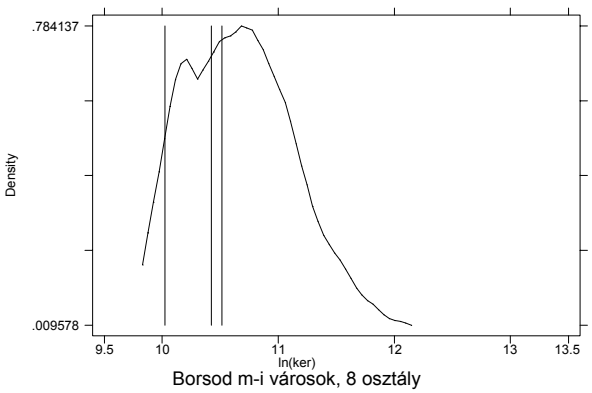
[12] Wage distributions stem from two subsequent surveys (from 1994 and 2001). Samples covered workers who entered employment from the UI register within the span of a month in 1994 and 2001. Wages are subjective estimates of the entrants as their own maximum and minimum expected wages. Wages underlying the distribution were calculated as averages of these expected minima and maxima

APPENDIX 4

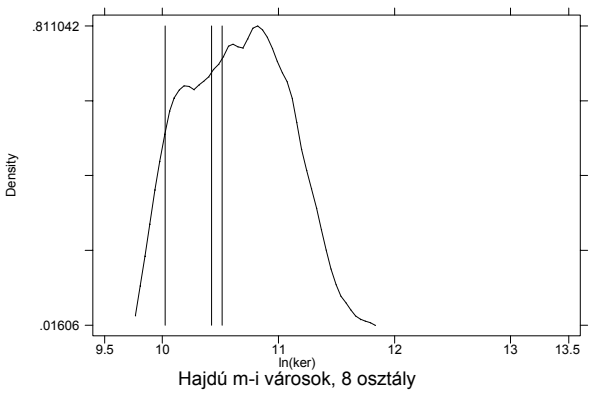
THE DISTRIBUTION OF LOG WAGES OF WORKERS WITH COMPLETED OR INCOMPLETE PRIMARY SCHOOL IN THE NON-BUDGET SECTOR, 1999 (THE DISTRIBUTIONS ARE SPLIT BY THE SUBSEQUENT MINIMUM WAGE LEVELS; MINIMUM WAGES ARE CALCULATED AT 1999 REAL VALUE)

Some regions where the proportion of unskilled workers is very high



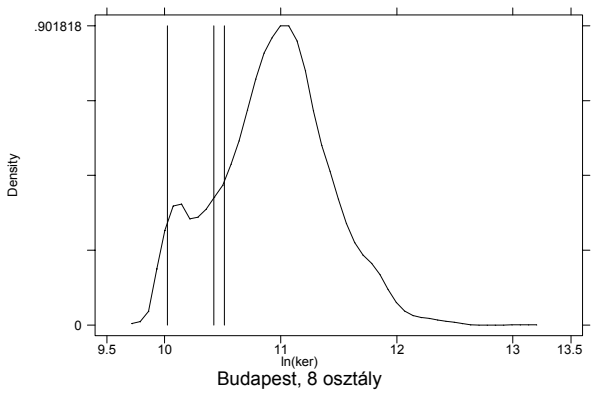


Borsod m-i városok, 8 osztály

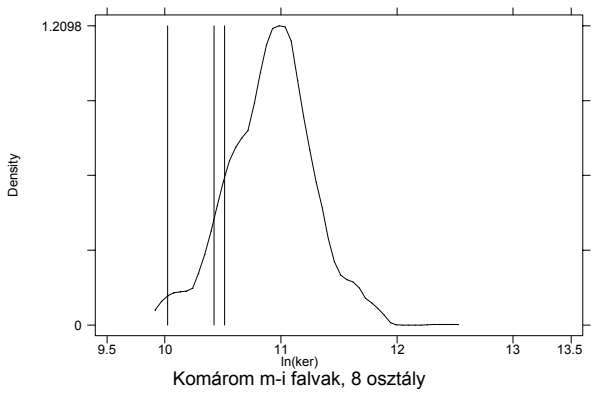


Hajdú m-i városok, 8 osztály

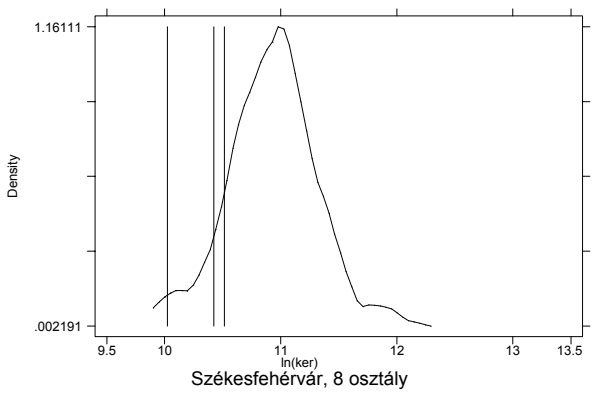
Some regions where the proportion of unskilled workers is very low



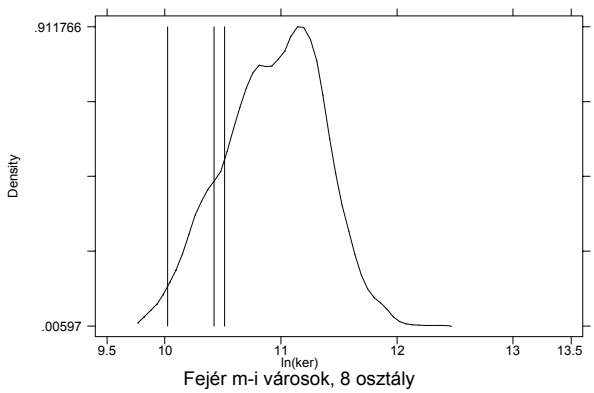
Budapest, 8 osztály



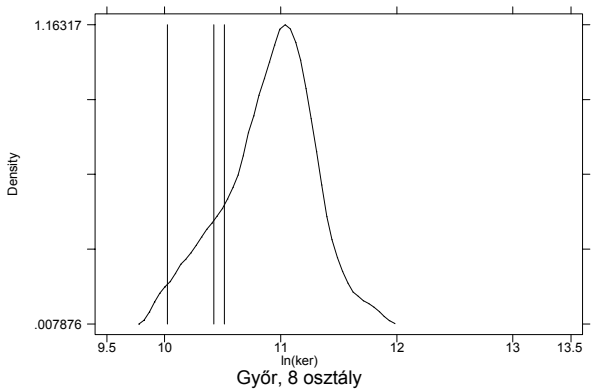
Komárom m-i falvak, 8 osztály



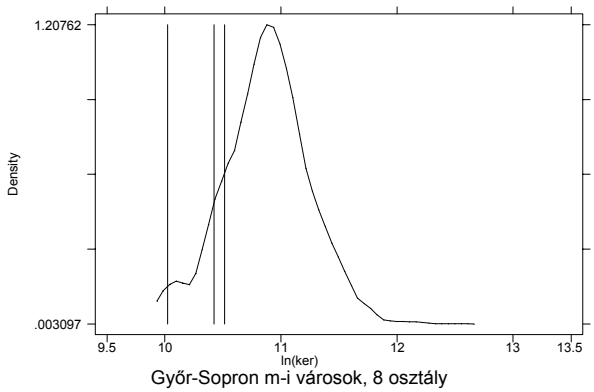
Székesfehérvár, 8 osztály



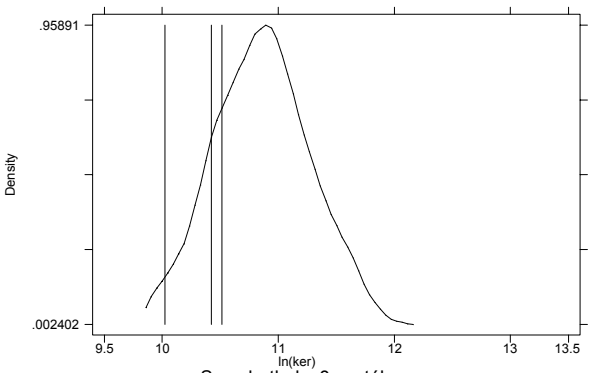
Fejér m-i városok, 8 osztály



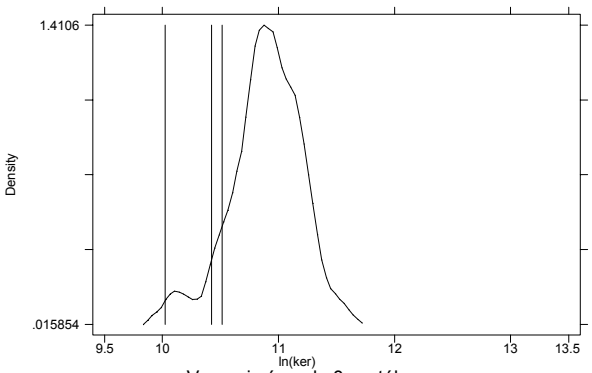
Győr, 8 osztály



Győr-Sopron m-i városok, 8 osztály



Szombathely, 8 osztály



Vas m-i városok, 8 osztály