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# Foreign Direct Investment, Wage Inequality, and Skilled Labor Demand in EU Accession Countries

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**Abstract**

During the 1990s Poland, Hungary and the Czech Republic have experienced rapid increases in wage inequality between skilled and unskilled workers and received the largest FDI inflow in Central and Eastern Europe. This paper analyzes whether FDI has contributed to the raise in earning inequality via a change in the skill composition of labor demand in the three countries. While we find that in Hungary and the Czech Republic FDI exerts a positive direct impact on the skill-premium, in none of the countries considered FDI has worsened wage inequality by favoring labor demand shifts.

J*EL classification*: F16, F23, J23, J31
*Keywords*: Foreign direct investment; Labor demand; Wage inequality.

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## 1. Introduction

This paper studies empirically the marked increase in the wage differentials between skilled and unskilled workers experienced by Poland, Hungary and the Czech Republic during the transition process to a market economy.

The question we ask is whether it is possible to explain the pattern of growing wage inequality in these EU accession countries using arguments based on increasing international integration, as those advanced first for developed and subsequently also for developing countries. In particular, motivated by the recent huge inflow of Foreign Direct Investment (FDI) in Poland, Hungary and the Czech Republic, we ask whether foreign capital penetration plays a role in explaining the pattern of wage inequalities in these countries via an expansion in the relative demand for skilled labor.

A marked increase in the wage differential between skilled and unskilled workers has been experienced during the 1980s and the 1990s in the United States and other industrialized countries. There is no agreement about the possible explanations of this pattern. Some studies focused on the skilled-biased technical change (SBTC), caused by the spreading utilization of computers and other new technologies within the firms (among others, Berman et al., 1994; Autor et al., 1998). Others focused on factors linked to increasing international integration. Within this explanation, three main sources of wage inequality have been proposed: first, trade in final goods and competition from low-wage countries; second, trade in intermediate inputs; finally, FDI and the activities of multinational enterprises (MNEs). The theoretical rationale for expecting an effect of trade in final goods on wage inequality is based on the well-known Stolper-Samuelson theorem, while the cross-border fragmentation of the production process and the de-locatization of stages of production in low-wage countries are expected to act as within-industry demand shifter toward skilled workers, that is the same effect of skill-biased technical change.[1] The role of FDI and multinationals in explaining raising wage inequality has received specific attention. Generally speaking, MNEs are likely to be more skilled labor-intensive than national firms and thus relative demand for skilled labor tends to rise as multinationals emerge and substitute for national enterprises (Markusen and Venables, 1997). Moreover, FDI is a channel through which fragmentation of production can take place. Firms in the skill labor-abundant (developed) country are likely to outsource the least skill-intensive activities, raising the skill-intensity of production and contributing to increase relative demand and relative earnings of the higher skilled in the home economy (Feenstra and Hanson, 1996).[2]

The rise in wage inequality was observed not only in developed countries, but also in many developing economies. This pattern conflicts with the neoclassical explanations about the redistributive effects of international trade. From a different theoretical perspective, Feenstra and Hanson (1996) focussed on the role played by foreign capital inflows. They argued that FDI and outsourcing by MNEs can raise relative demand for skilled labor not only in the home (industrialized) country but also in the host (developing) one, increasing wage inequality in both economies. They found that capital transfer by US firms can account for the bulk of the increase in the skilled labor share of total wage bill that occurred in Mexico in the late 1980s (Feenstra and Hanson, 1997). Despite a growing empirical literature investigating the possible determinants of wage inequality in LDCs, the results are mixed and the sources of this pattern

[1] For a survey on the existing theoretical and empirical literature on international outsourcing, see Feenstra and Hanson (2001).

[2] There are only few empirical studies on this issue and their results show no systematic correlation between raising wage inequality and either outward or inward FDI in developed countries. See Slaughter (2000) and Blonigen and Slaughter (2001) on US, and Head and Ries (2002) on Japan.

remain a puzzle.[3]

During the process of transformation from planned to market economies, also Central and Eastern European Countries (CEECs) have experienced a substantial increase of inequality in labor earnings. The CEECs are, in some sense, in the middle between developed and developing countries. While relatively well endowed of human capital, as a result of the large access to education of the centralized regimes, these countries nonetheless suffer a mismatching between their current institutional, economic and political system and the requirements posed by a well functioning market economy. It is therefore important to investigate whether the huge inflow of foreign capital has played a role by changing the relative labor demand as suggested for developed and developing countries, or other factors, such as institutional changes and increasing trade integration with the EU countries, might help to explain recent labor market outcomes. To answer this question, we estimate the impact of increasing inflow of FDI on wage inequality and on the composition of labor demand between skilled and unskilled workers in Poland, Hungary and the Czech Republic.

Our main finding is the following. While in Hungary and the Czech Republic FDI exerts a positive direct impact on the skill-premium, in none of the countries considered FDI has worsened wage inequality by favoring labor demand shifts. In the case of Poland it is difficult to trace any specific role at all for FDI, which may be due to the slower pace of the transition to a market economy that characterizes this country. Thus, the emerging picture is one in which multinational firms play an active role in the transition-induced restructuring process, but mainly by favoring workers mobility across sectors and occupations and bringing wage dispersion closer

[3] Most of the empirical studies on the relationship between international integration and wage inequality in LDCs has analyzed the effects of trade liberalization episodes. For a recent survey, see Goldberg and Pavcnik (2004).

to the actual skill distribution, rather than influencing the labor demand composition.

The structure of the paper is as follows. The next section presents the dataset used in the empirical analysis and the descriptive evidence on the three accession countries under analysis. Section 3 sets up the empirical model. Estimation and empirical results are shown and discussed in Section 4. Section 5 concludes.

## 2. Data description and stylized facts

Our analysis focuses on the manufacturing sector of Poland, Hungary and the Czech Republic and aims at investigating whether and to what extent the huge inflow of FDI received by these countries in the last decade might have contributed to the raising wage inequality between skilled and unskilled workers which has been experienced since the beginning of the transition process. At this purpose, lacking a more detailed industrial breakdown on inward FDI, we use an industry-year panel consisting of observations on six sectors for the period 1993-2000. To proxy MNEs penetration, we use the inward FDI position (stocks), primarily relying upon the OECD "International Direct Investment Statistics" database. Earning inequality and composition of labor demand are proxied by the skilled labor shares of total wage-bill and employment, respectively. We follow the usual skill approximation and accordingly define non-manual and manual employees as skilled and unskilled workers, gathering data on employment and wages for the two categories from each country's statistical yearbooks. To take into account other possible and often debated explanations for the rise in wage inequality, our database includes also the following variables: business enterprise expenditure on R&D, proxing the influence of technical progress; exports and imports of final goods, accounting for the effects of international trade on relative factor prices; gross value added, controlling for industry-scale effects. All data are converted in constant 1995 US$ and aggregated according to the FDI classification. For a detailed description of the database, see the Appendix.

Some stylized facts on these data are as follows. Poland, Hungary and the Czech Republic have experienced rapid increases in the wage differential during the 1990s: between 1993 and 2001, the relative wage of skilled workers has increased from 1.40 to 1.81 in the Czech Republic and from 1.94 to 2.29 in Hungary, whereas it has risen from 1.44 to 1.80 in Poland between 1994 and 2001. At the same time, relative employment of skilled workers – as proxied by the ratio between the number of non-manual and manual employees - has shown a moderate decline in the case of Hungary and the Czech Republic (from 0.29 to 0.25 and from 0.38 to 0.33, respectively); in Poland, instead, it has slightly increased since 1994 (0.28) and today amounts to 0.31 (Figure 1). The skilled labor shares of total manufacturing wage bill and of total employment have moved consistently with these patterns.[4] These developments are somewhat conflicting with the evidence from other developed and developing countries, where higher wage inequality has been generally caused by a growing demand for skills. At a first glance, this suggests that the opening of the wage gap might not have been induced –at least in Hungary and in the Czech Republic- by any of the several forces which usually worsen wage disparities by raising the relative demand for skilled labor. An additional piece of empirical evidence may provide useful insights. Decomposing the changes in the skilled labor shares of total employment and of total wage bill, Crinò (2004) finds that they are mostly due to within-industry rather than

[4]Looking at the dynamics of relative wage and employment at an industry level, one common feature for the three countries emerges: the highest increase in earning inequality has been experienced in the food and chemical industries. The evolution of the measures of wage and employment disparities appears much more differentiated in the remaining sectors (Crinò, 2004).

beetwen-industries variations. Therefore, whatever the causes of wage inequality, they seem to act through changes in the proportion of non-manual workers within each industry, rather than through the reallocation of the workforce towards industries with different skill-intensity. This is in line with findings from previous empirical analysis on different DCs and LDCs countries (Table 1).

Over the same period, Poland, Hungary and the Czech Republic have received the largest FDI inflow in Central and Eastern Europe; along with Slovakia and the Russian Federation, they account today for 3/4 of the region's inflows. In percentage of GDP, FDI stocks have sharply increased in the last decade raising from values below 4% in 1990 to more than 40% in Hungary and in the Czech Republic and than 20% in Poland in 2000 (UNCTAD, 2002). Consistently, FDI stocks have markedly increased also in the manufacturing sector, which accounted for nearly half of all inward stock, and some interesting regularities emerged in terms of sectoral distribution. At the beginning of the transition process, foreign capital was concentrated primarily in the food sector (accounting for more than 35% of total manufacturing FDI stock), and, to a lesser extent, in textile and wood activities (Poland and Hungary), non-metallic products (Czech Republic), and chemical products (Poland and Hungary). During the 1990s, with the general huge increase of FDI inflows, foreign investors have been progressively redirecting their investments towards more capital and skill-intensive industries. Nowadays, more than 50% of total manufacturing FDI stock is concentrated in three industries: chemicals, metal and mechanical products, transport equipment. In Hungary and Poland, the food industry still accounts for a share around 30%.[5]

FDI is just one expression of the rapid process of international integration which the three

[5] For a more detailed industry-level analysis, see Crinò (2004).

countries have been involved in after the fall of the communist regime in 1989. Such a process, in fact, has also resulted in a rapid increase and in a marked reorientation of trade flows towards new partners, especially the more advanced EU countries.[6] Likewise, real exports and imports have almost doubled in manufacturing and their industrial composition has changed. Starting with a trade pattern typical of less developed countries, with high shares of exports in textiles, food (especially Hungary), and metallurgical industries (Czech Republic and Poland), the three countries have experienced important changes in their comparative advantages. Progressively, the pattern of specialization has shifted towards skilled labor intensive and capital intensive sectors. The contribution to total manufacturing exports of the machinery, computers and communications and of the transport equipment industries has increased dramatically, reaching a share of more than 40% in Hungary and the Czech Republic (only 33% in Poland). Together with these developments in inter-industry specialization, a general increase in the relevance of intra-industry trade has occurred, particularly in trade relations with the EU. This is the result, on one hand, of the growing importance of global production networks and the role of MNEs in integrating these countries into the international division of labor, and on the other hand, of the improvements obtained in intra-branch product quality (Landesmann and Stehrer, 2002).

In conclusion, since the transition process started in 1989, the three countries have remarkably opened up to international integration, allowing higher penetration by foreign investors and intensifying their participation in trade flows. At the same time, earning inequality in manufacturing has worsened and skilled labor has become more rewarded. The contemporary occurrence of these phenomena, however, does not necessarily imply a causal relationship between them.

[6] The EU-15 accounts today for about 70% of total exports and imports, while no more than 50% of total flows were directed towards this area in 1990.

The remaining part of the paper, therefore, investigates the existence of such a causal link, by making use of nonparametric and parametric approaches.

## 3. The empirical model

Empirical studies in international trade on the determinants of skill upgrading and skill premium typically maintain efficient allocation of resources within industries, either explicitly or implicitly. We put ourselves into this tradition, and for our empirical analysis we maintain cost minimization within sectors for given stock of capital. In more detail, we assume the existence of a representative firm for a given industry, which minimizes the cost of skilled and unskilled labor to produce a given amount of output, treating capital as fixed over the relevant sample period. This optimization problem for the firm yields the following variable cost function

$$C\left(w^s, w^u, Y, K\right) = \min_{N^s, N^u} \left[w^s N^s + w^u N^u : Y = F\left(N^s, N^u, K\right)\right] \tag{3.1}$$

where $Y$ denotes output; $K$ denotes capital; $N^s$ and $N^u$ denote the number of skilled and unskilled workers; and $w^s$ and $w^u$ denote wages for skilled and unskilled workers.

One clear advantage of cost minimization is the possibility of testing structural characteristics of interest, such as returns to scale and elasticities of substitution, from estimated demand or share equations. It is, however, questionable that in Poland, Hungary and the Czech Republic instantaneous adjustments of inputs to relative price variations occurred without exceptions of sectors and years, especially if one thinks of the dramatic structural changes that these countries have experienced over the period considered. For this reason, before estimating parametric share equations based on cost minimization, we first test cost minimizing behavior by direct inspection of data. This assumption, indeed, can be tested nonparametrically by verifying the inequalities implied by the Weak Axiom of Cost Minimization (WACM) (see Varian, 1984) on sector data:

$$Y_t \leq Y_r \text{ and } K_t \geq K_r \Rightarrow w_t^s N_t^s + w_t^u N_t^u \leq w_t^s N_r^s + w_t^u N_r^u, \; \forall t, r = 1, ..., T \tag{3.2}$$

where $t = 1, ..., T$ indicates the year, so that subscription denotes an observation for the corresponding variable. Given the limits of our data, some modifications to (3.2) are however necessary. Since we do not observe the capital stock, we assume that it remains fixed over any two contiguous years. In addition, we allow for non-regressive technical progress by restricting comparisons of the current cost levels to the past year only, so that the inequalities in (3.2) modifies as follows

$$Y_t \leq Y_{t-1} \;\; \Rightarrow w_t^s N_t^s + w_t^u N_t^u \leq w_t^s N_{t-1}^s + w_t^u N_{t-1}^u, \; \forall t = 1, ..., T. \tag{3.3}$$

The evidence drawn from the nonparametric analysis will be used for searching adequate econometric specifications and estimators for the cost function. The baseline parametric model is derived by the following translog function for (3.1):

$$\begin{aligned} \ln(VC) \quad = \quad & \beta_0 + \sum_{i=s,u} \beta_i \ln w^i + \beta_Y \ln Y + \beta_K \ln K \\ & + \frac{1}{2} \left( \sum_{i=s,u} \beta_{ii} \ln^2 w^i + \beta_{YY} \ln^2 Y + \beta_{KK} \ln^2 K \right) + \\ & + \beta_{su} \ln w^s \ln w^u + \sum_{i=s,u} \beta_{Yi} \ln w^i \ln Y + \sum_{i=s,u} \beta_{Ki} \ln w^i \ln K \end{aligned} \tag{3.4}$$

By applying Shephard Lemma and exploiting the homogeneity and adding-up restrictions for the translog, we can derive the following share equation for skilled labor:

$$WSh_s = \beta_s + \beta_{ss} \ln \left( \frac{w^s}{w^u} \right) + \beta_{Ks} \ln K + \beta_{Ys} \ln Y \tag{3.5}$$

where $WSh_s$ is the skilled-labor share of total industry wage-bill. Modifications of equation (3.5) will be estimated and tested using a battery of estimators on the pooled data for the three countries, but also on subsets of data.

## 4. Results

The first part of this section contains results for the nonparametric analysis. The rest of the section is devoted to parametric results.

### 4.1. Nonparametric analysis

From the analysis of WACM as in (3.3) we find the following (Table 2). For the Czech Republic there are 15% rejections overall; 7% rejections greater than 10%; and 4% rejections greater than 20%. There is only one critical year, 1997, with significant rejections (>5%) in sectors Wood, paper and publishing, and Basic metals and metal products. For Hungary there are 28% rejections; with 17% rejections greater than 5%; and 11% rejections greater than 20%. Critical years are 1999 and 2001, with significant rejections in sectors Food; Chemicals; Metal and mechanical products; and Transport equipment. Poland is the country that seems the most distant from sectoral cost minimization, with 45% rejections of WACM. 36% rejections are greater than 10% of which 27% are greater than 20%. Violations, however, are concentrated in

1995 and 1996, and are particularly significant in sectors Food; Textiles; Chemicals; Metal and mechanical products.

To sum up, for no country we are able to observe systematic violations of WACM spread over all sectors and years. Instead, we do observe for all countries a few critical years and sectors, for which deviations from cost minimizing behavior emerge as particularly severe. We shall take these findings into account for parametric estimation in the next section, when relaxing somewhat the homogeneity of the specification over time.

### 4.2. Parametric analysis

By simply adding extra regressors, equation (3.5) can be expanded to take into account the effects of several variables on wage inequality. First of all, we use inward-FDI stocks, $(FDI)$, as measure of multinational enterprises penetration. Moreover, in order to take into account the other variables that according to the existing literature might exert effects on the wage gap between skilled and unskilled workers (Berman et al., 1994; Autor et al., 1998; Machin and Van Reenen, 1998; Egger and Stehrer, 2003) we control also for the influence of international integration by using exports and imports of final good, $(X)$ and $(M)$, and proxy technological change with the total business enterprise expenditure on R&D, $(R\&D)$. Industry output, $Y$, is measured by gross value added, $(VA)$. A set of year-dummies, $(Year)$, enters the regression to capture time-specific, industry-invariant shocks which are likely to be important in these countries because of the deep institutional changes brought about by transition. Since there are no observations available on the capital stock at an industry level, we identify our share equation by assuming that all variations of capital stock correlated with the included explanatory variables are captured by sector- and time-dummies. Finally, it is reasonable to assume that neither FDI nor any of the other variables exert an immediate impact on relative employment and wages; all regressors, therefore, enter the regression lagged once. This gives us the following baseline estimating share equation:

$$\begin{aligned} WSh_{i,c,t} &= \alpha_{i,c} + \beta_c^w \ln\left(\frac{w^s}{w^u}\right)_{i,c,t-1} + \beta_1 \ln(FDI)_{i,c,t-1} + \beta_2 \ln(X)_{i,c,t-1} + \\ &\quad + \beta_3 \ln(M)_{i,c,t-1} + \beta_4 \ln(VA)_{i,c,t-1} + \beta_5 \ln(R\&D)_{i,c,t-1} + \gamma(Year)_t + \varepsilon_{i,c,t} \end{aligned} \quad (4.1)$$

where $\varepsilon$ is an idiosyncratic error term and $\alpha$ is an industry-specific fixed effect; $i = 1, ..., 6$ indexes the manufacturing industries, $t = 1993, ...2000$ is the time span of our panel[7] and $c = PO, HU, CR$ are the three countries considered; the subscript of the coefficient on the log-relative wage of the skilled, $\beta_c^w$, indicates that the corresponding variable is interacted with the country-dummies. Constant returns to scale in production imply $\beta_4 = -(\beta_1 + \beta_2 + \beta_3 + \beta_5)$.

#### 4.2.1. Fixed-effect estimation

Several studies on the skill upgrading (Machin and Van Reenen, 1998; Blonigen and Slaughter, 2001; Pavcnik, 2003) have noted that the log-relative wage of the skilled is likely to be endogenous in a share equation. In fact, as suggested in Berman et al. (1994), although part of the variation in the relative wage can be due to differences in the skill-mix across industries, some of it is likely to be caused by the skill upgrading occurring within industries. To avoid this problem, the relative wage is usually excluded from the regression and its effect is taken into account by using time- or country-dummies. Following this line of reasoning, we first omit such variable and

[7] In the case of Poland, the time span is 1994-2000.

estimate equation (4.1) by a fixed-effect estimator allowing for unobserved heterogeneity across industries; results are presented in Table 3. To deal with the small cross-sectional dimension of our panel, we begin pooling the three countries together and accommodating the influence of country-specific characteristics by diversifying the industry fixed effects across countries; as a result, 138 observations become available (column 1).[8] The fixed effects are always highly significant. No evidence does emerge of an impact of FDI on the skilled-labor share of total wage-bill; although positive, the coefficient on FDI is very small and not significantly different from zero. On the other hand, our export measure is highly significant and its negative coefficient suggests that trade integration might contribute to a reduction in the relative demand for skilled labor. Jointly significance of time dummies, instead, reveals that industry-wise year-specific changes -implied by the ongoing transition process- have hit the manufacturing sector of the three countries and exerted important effects on the relative demand for skilled labor. Such effects, of course, are likely to differ across the three countries; it is, indeed, realistic to think of such year-specific shocks as affecting the relative demand of skilled labor differently from one country to the other, because of differences in the pace and speed of the transition to a market economy. To allow for country specificities, we reestimate equation (4.1), interacting the time-dummies with the country-dummies (column 2). The results seem very robust. In particular, no relationship emerges between foreign presence and the relative demand for skilled labor: not only the coefficient on FDI is still not significant, but it switches sign and becomes negative as the specification is expanded by including the interacted time-dummies. As expected, the latter are highly significant and support the intuition about strong heterogeneity in the year-specific

[8] More specifically, our panel consists of 6 manufacturing industries for 8 years (1993-2000) in the cases of Hungary and the Czech Republic, and 6 industries for 7 years (1994-2000) in the case of Poland.

shocks across the three economies.

#### 4.2.2. Instrumental Variable estimation on the pooled sample

So far, we have dealt with the log-relative wage endogeneity by simply omitting such variable from the regression. Microeconomic theory, however, implies that cost minimizing relative factor-demands are functions of relative prices and would therefore suggests to keep into explicit consideration the log-relative wage of the skilled when estimating a share equation like (4.1). This task, however, requires to change estimation method in order to tackle the endogeneity problem. Finding valid instruments for relative prices would allow to include them explicitly in the specification dealing at the same time with the endogeneity. We, therefore, estimate again equation (4.1) with an instrumental variable (IV) estimator; we choose as instruments for the country-specific relative skilled wages their second and third lags. In this setup, moreover, it is worth nothing that an endogeneity problem might indeed affect not only relative prices but also the FDI variable. In fact, there might be, first of all, some common forces jointly determining both the skilled-labor share of total wage-bill and the foreign penetration in a country; furthermore, as noted in Pavcnik (2003), it might well be the case that in some industries it is foreign penetration to be incentivated by an already high skill-intensity, rather than the opposite. For this reason we instrument also our FDI variable by using its own first three lags as well as the first three lags of a country-specific FDI measure, obtained by interacting $FDI$ with the country-dummies[9]. Results are reported in Table 4. In each specification, the Sargan's test does not reject the null hypothesis of instruments validity at any conventional significance

[9] By the same token, exports and imports might be endogenous too. We have instrumented them using a similar instrument set. The results, however, do not change at all when doing this. All these regressions are available from the authors upon request.

level; the constant returns to scale hypothesis is instead rejected. When relative wages enter the regression, time dummies loose their significance (column 1); this holds true also when they are interacted with the country-dummies (column 2). This suggests that the previously found time-effects on the relative skilled-labor demand act through changes in the relative price of the skilled-labor input. Indeed, after excluding the time-dummies from the set of regressors[10], relative wages become highly significant (column 3). Given their positive sign, the matrix of estimated coefficients on relative prices is not negative semidefinite. This not withstanding, the matrix of substitution elasticities may be negative semidefinite, as required by cost minimization. For this to be true a necessary condition is that the elements on the main diagonal must be negative, that is $\varepsilon^c_{s,s} = \frac{\beta^w_c + \overline{WSh^2_c} - \overline{WSh_c}}{\overline{WSh^2_c}} < 0$ for $c = PO, HU, CR$; since this formula assumes no optimization error, it has to be computed using fitted, rather than actual, average shares of skilled labor in total costs (Hayashi, 2000). As evident, reported shares imply that the regularity condition is satisfied for both Hungary and the Czech Republic but not for Poland. Turning to the other variables, these regressions further strenghten the results from the fixed-effect estimation: FDI does not exert any significant influence on the relative demand for skilled labor and, indeed, estimates for the associated coefficients are rather small and actually characterized by the negative sign; on the contrary, further evidence is provided for the influence of exports as acting in favor of the unskilled and against the skilled.

Hence, regardless of the huge and contemporary increase in foreign penetration and in wage inequality, our results do not support the idea of a positive causal relationship between the two phenomena. FDI does not seem to have affected relative labor demand for the skilled and thus cannot be taken as responsible for the deepening in the wage gap through outward

[10] Dummies are still included, however, in the instruments matrix.

shifts of the relative demand schedule. We have indeed showed that relative employment of the skilled has actually fallen during the last decade, at least in Hungary and in the Czech Republic; increasing export flows seem to be responsible for some part of this reduction and accordingly to have contributed to containing the documented worsening in wage inequality. Nevertheless, the past socialist experience of these three countries suggests that the rising in earning differentials may have occurred even lacking a biased labor demand shift in favor of the skilled. To be more precise, all formerly centrally planned economies were certainly well endowed of human capital, as evident from comparisons with other economies either at the same or at higher level of development; at the same time, they showed a very compressed wage structure unable to reflect the actual skill distribution. In these respects, the transition to the market has brought about two not negligible changes. On the one hand, wage determination has gained flexibility and approached schemes closer to those typical of a market economy; as a result, the earning structure has changed and progressively fitted much better the actual skill distribution. A number of studies documents, indeed, significant rises in the returns to education during the 1990s as compared to the period under the communist regime.[11] As a result, wage inequality has been physiologically worsened by the transition process. On the other hand, skills available before the transition could not be easily adapted and transferred in the developing market economy, because they were too specialized and specific for a given firm or production process (Boeri, 2000; Campos and Coricelli, 2002; Commander and Kollo, 2004; Svejnar, 1999); this has caused, on one side, a strong reduction in secondary schools and vocational-school enrollment rates (Boeri, 2000; Campos and Coricelli, 2002) and, on the other side, increasing unemployment

[11] See Boeri and Terrell (2002), Fleisher et al. (2004), and Svejnar (1999), for a comprehensive review of relevant studies.

for workers with vocational education (Boeri and Terrell, 2002; Flanagan, 1998). These two transition-induced labor market changes explain the documented peculiar pattern of relative employment and wages in Hungary and in the Czech Republic: wage inequality increased as a consequence of the departure from the highly compressed wage structure in place under the socialist regime and skilled labor became progressively more rewarded. At the same time, the needs of the newborn market economy required to adopt efficiency criteria in resource allocation and to rationalize the use of inputs in production, included the overspecialized pre-transition skills which became, as a result, needless and unadaptable in the new private sector. Among the three countries, Hungary and the Czech Republic were interested from these changes sooner than Poland, which lagged behind in the microeconomic and industrial restructuring in spite of the rapid completion of the macroeconomic aspects of transition. The role played by foreign firms in the restructuring process and their contribution to the above changes is undeniable if we consider that multinationals corporations have penetrated these economies by mainly exploiting the opportunities opened up by the ongoing privatization process, so helping creating a new private sector and exposing it to stronger competitive pressures. Through this channel, FDI might have worsened wage inequality by simply contributing to change the wage determination and to increase the skill-premium with respect to the pre-transition period. This intuition is indeed supported by the results reported in column 4 of Table 4, where we use the skill-premium -as proxied by the ratio between skilled and unskilled wages- as dependent variable, rather than the share of skilled labor in total wage-bill: FDI is highly significant and positively signed; at the same time, the joint significance of time-dummies reveals that institutional changes and the above mentioned features of the transition process have exerted a crucial impact on the dynamics of the wage premium in each country during the 1990s.

### 4.2.3. Instrumental Variable estimation on the country-specific subsamples

Up to now, we have jointly analyzed the three economies. Although this allows to benefit from more degrees of freedom, it might also distort the results for two reasons. First, the stylized facts on the labor market have shown a very different behavior of the relative employment in Hungary and in the Czech Republic, on the one hand, and in Poland, on the other. Second, both the nonparametric and the parametric analysis has questioned the validity of a cost minimization setting for analyzing the dynamics of wage inequality in Poland. In order to check the robustness of our results, we therefore proceed by repeating the instrumental variable estimation on two subsamples: the first includes Hungary and the Czech Republic, whereas the second consists of observations for Poland only. We begin presenting estimates from the first subsample. As far as the share equations are concerned, the first three columns of Table 5 show that the same conclusions drawn before apply to the case of Hungary and the Czech Republic: the constant returns to scale hypothesis is rejected; time dummies (whether interacted or not with country-dummies) are not jointly significant and excluding them from the regression enhance significance in the country-specific wage variable; regularity conditions for the matrix of substitution elasticities are satisfied at reported average fitted shares. FDI does not seem to exert any significant impact on wage inequality through upward shifts in the relative demand for skilled workers and indeed, significance aside, its associated coefficient is again negative in all specifications. Still, FDI cannot be denied at all an influence on the increased wage inequality occurred in the two countries in the 1990s; as above, in fact, estimating a wage equation yields a significant and positive impact of foreign presence on the skill-premium as column 4 of Table 5 shows; we attribute such an impact to the contribution of FDI to the transition process towards the market economy.

The above results, although robust for Hungary and the Czech Republic, do not seem to be confirmed in the case of Poland. Before presenting them, one estimation issue should be emphasized and taken into account from now on. Namely, the IV estimator has asympthotic properties which require a large number of observations; when concentrating on the Poland case, the dimension of our panel dramatically shrinks and we are left with only 42 observations; with this *caveat* in mind, however, we still believe it is worth trying to study this country separately for the reasons illustrated before. And indeed, interesting insights do emerge from this analysis. First of all, share equation estimates reported in the first two columns of Table 6, yields very different results from those related to the other two countries: apart from the relative wage, whose significant and positive coefficient in the specification without time-dummies violates the regularity conditions for the substitution elasticity matrix, no other variable is significantly able to explain the skilled-labor share of total wage-bill. Consistently with what we have seen for the other two countries, the effects of institutional changes on the wage premium do emerge from the wage equation reported in the third column of Table 6, as the time-dummies are jointly significant; if any, however, the effect of FDI on relative wages seems to have been negative and this would sharply conflict with what has happened in the other two cases. Before drawing any conclusion from these results, one point should be stressed. The cost minimization setup which we are working in seems to be inadequate to describe the Polish case; on the one hand, in fact, the nonparametric analysis has revealed that the worst violations of the WACM have occurred in this country; on the other hand, the regularity conditions for the substitution elasticity matrix were never satisfied in any of the reported specifications. For these reasons, we now try to exploit the findings from the nonparametric analysis in our econometric estimation; namely, we eliminate from the sample those years which have proved to be the most problematic in the light of the cost minimization model: 1995 and 1996. Results are shown in the last two columns of Table 6. Our intuition is confirmed; the coefficient on the log-relative wage in the share equation, in fact, is now insignificant, which is enough to guarantee a negative semidefinite Hessian of the cost function; although positively signed, the FDI coefficient is not significantly different from zero, suggesting once again that foreign presence is not responsible for the rise in wage inequality either in Poland, at least through its effects on the relative demand for skilled labor; export activity, instead, acts significantly on the relative labor demand in favor of the unskilled portion of the workforce; institutional changes and the rising speed of firm restructuring, finally, play a significant role in Poland, differently from what we have learned for the other two countries. This last result has to be attributed to the slower pace which has characterized Poland in the completion of the microeconomic aspects of transition and which has implied important renewals and restructuring in the manufacturing sector of the country also during the last decade. Finally, as opposed to the other two countries, the FDI now does not seem to exert a significant impact on relative wage either, which confirms the specificity of this country.

## 5. Conclusion

This paper investigates whether FDI penetration has contributed to the raise in wage inequality and to the change in the skill composition of labor demand in three EU accession countries - Poland, Hungary and the Czech Republic - during the 1990s. Using nonparametric and parametric approaches our results suggest that, although FDI has not worsened wage inequality by favoring labor demand shifts in the three countries, it has contributed to raising the skill-premium through the active role played by multinational firms in the transition-induced restructuring process in Hungary and the Czech Republic. In more detail, for these two countries the transition and privatization process, hand in hand with an increasing international integration, has pushed the labor market from a compressed and rigid wage structure towards forms of wage determination more typical of a decentralized market economy. Overall, the FDI penetration more than affecting production technology and relative labor demand seems to have favored workers mobility across sectors and occupations and brought wage dispersion closer to the actual skill distribution. For Poland, while the scarce effect of FDI on the relative skilled labor demand is confirmed, evidence does not support a significant direct effect of FDI on the skill-premium, so that it is difficult for us to carry over the argument advanced for Hungary and the Czech Republic into this case. The impossibility to tell a common story for the three countries can be explained by the slower pace of the transition process taking place in Poland.

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## DATA APPENDIX

Data on Foreign Direct Investment stocks in Poland and Hungary come from the OECD "International Direct Investment Statistics" (http://www.sourceoecd.org) and are classified in six ISIC Rev.3 industries. Unfortunately, the OECD data offer a sufficiently long time series only for Poland (1994-2000); for Hungary, instead, they cover only the period 1993-1998. In order to obtain a longer time series (1993-2000), therefore, we merge them with the data provided for more recent years by the UNCTAD "World Investment Directory" (http://www.unctad.org), whose observations for 1997 and 1998 fully coincide with the OECD ones. For the Czech Republic, we rely upon the "Balance of Payments Statistics" of the Czech National Bank (www.cnb.cz), which provide data on FDI inflows in eight manufacturing industries for the period 1993-2001.[12] FDI stocks have been obtained as cumulative inflows.

Data on exports, imports, and gross value added are collected from the OECD "Stan Database for Industrial Analysis".[13] Data on total business expenditure on R&D come from the OECD "Basic Science and Technology Statistics" (http://www.sourceoecd.org). All data are based on the ISIC Rev.3 classification.

Finally, we refer to each country's Statistical Yearbook for information on wages and employment of the manual and non-manual workers. Such data are classified according to NACE Rev. 2 in fourteen 2-digit manufacturing industries and are available for the period 1993-2001. Some clarification is needed about the Czech labor market data. Statistical yearbooks for this country, do not report information on employment and wages for the non-manual workers. We therefore

[12] For the parametric analysis, the eight Czech manufacturing industries have been converted in the six ISIC Rev.3 industries.

[13] Observations on Polish gross value added in 2000 are gathered from the OECD "Structural Statistics for Industry and Services" (2003) (http://www.sourceoecd.org), whereas information on the corresponding Hungarian variable in 2001 comes from the same country's Statistical Yearbook.

estimate them following Egger and Stehrer (2003).[14] In particular, we first compute the number of non-manual employees as difference between total and manual employment; the non-manual workers wage is $w_i^s = N_i^s \left\{ w_i - [\left( \frac{N_i^u}{N_i} \right) w_i^u] \right\} / N_i^u$, where the superscripts index skilled (non-manual) and unskilled (manual) workers and the subscript the manufacturing industries, $N$ is the number of employees and $w$ the average wage.

All data are converted in constant 1995 US$. Deflation is carried out by using a common deflator across industries. It comes from EUROSTAT, NewCronos Database[15] and is defined as the ratio between GDP at current prices and GDP at constant prices.

As mentioned, the original data are classified according to very different industrial breakdowns; this requires us to make them consistent with the six-industry FDI classification, which clearly represents the most stringent constraint because of the limited industrial disaggregation. To recover the FDI classification, we carry out a two-step procedure. First, we convert the industries from ISIC Rev. 3 to NACE Rev. 2 (2-digit) classification according to the correspondence table provided by EUROSTAT (http://www.europa.eu.int/comm/eurostat), and then we aggregate the observations in the six ISIC Rev. 3 industries. The aggregation procedure for the wage variables needs to be explicitly explained; starting with data on manual and non-manual employment and wages at the NACE 2-digit level, we calculate the average wage of the two skill-categories in the six industries as follows: $w_A = \left( \sum_{i=1}^{I} w_i^j N_i^j \right) / \left( \sum_{i=1}^{I} N_i^j \right)$, where $j = s, u$ refers to the two skill-categories and $i = 1, ..., I$ indexes the NACE 2-digit sectors belonging to the ISIC Rev. 3 industry $A$; $\sum_{i=1}^{I} w_i^j N_i^j$ and $\sum_{i=1}^{I} N_i^j$ are, respectively, the total wage-bill and employment of the manual and non-manual workers in $A$.

[14] For the period 1993-1999, the data we use coincide with those in Egger and Stehrer (2003). We are grateful to R. Stehrer for the provision with data.

[15] On line at the following link: http://www.europa.eu.int/comm/eurostat.

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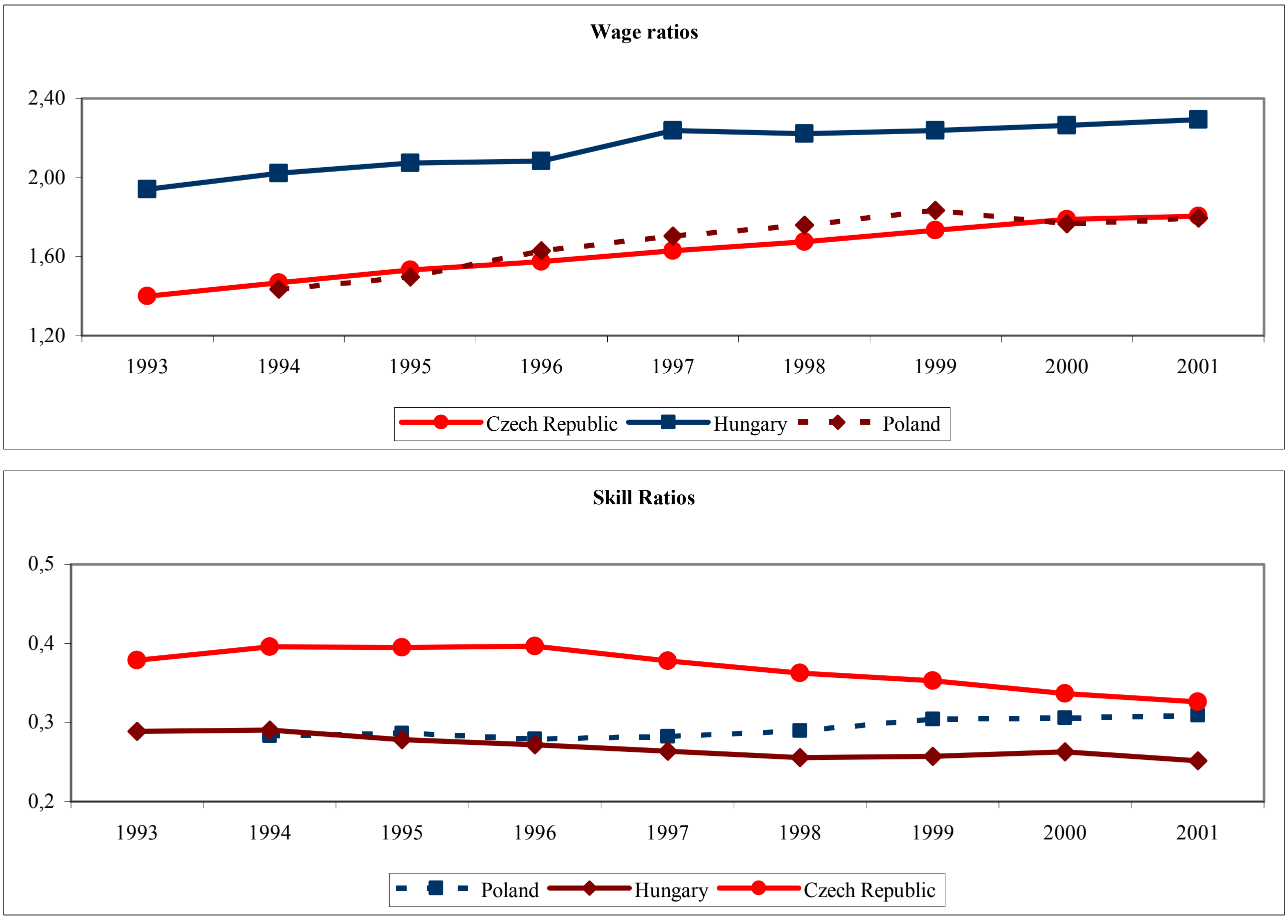
## Figure 1 – Skill and Wage Ratios in Poland, Hungary and the Czech Republic*




*Source: National Statistical Yearbooks (various years)*
*Skill ratio: non-manual employment relative to manual employment
Wage ratio: non-manual workers wage relative to manual workers wage

## Table 1 – Decomposition of the Change in the Skilled Labor Shares of Total Employment and Total Wage-bill (% values)*

| | Employment | | | Wage-bill | | |
|---|---|---|---|---|---|---|
| | Between | Within | Total | Between | Within | Total |
| **Poland** (1994-2001) | -0,065 | 1,564 | 1,49 | -3,37 | 8,28 | 4,91 |
| **Hungary** (1993-2001) | -0,74 | -1,50 | -2,24 | -0,47 | 1,51 | 1,04 |
| **Czech Republic** (1993-2001) | -0,80 | -2,67 | -3,47 | 0,60 | 2,40 | 3,00 |

*Source: Crinò (2004), Table 2.*

* Following Berman et al. (1994), the overall change in the skilled labor shares of total wage-bill and employment has been decomposed in the two dimensions, *between-industries* and *within-industry,* according to: $\Delta WSh_S = \sum_{i=1}^{I} \Delta P_i \overline{WSh_{Si}} + \sum_{i=1}^{I} \Delta WSh_{Si} \overline{P_i}$ for $i$ =1, .. $I$, where $WSh_S$ is the skilled-labor share of total wage-bill (employment) in manufacturing, $WSh_{Si}$ is the skilled-labor share of wage-bill (employment) in industry $i$, $P_i$ is the share of industry $i$ in total manufacturing wage-bill (employment), $\overline{WSh_{Si}}$ and $\overline{P_i}$ are time averages of the corresponding variables.

## Table 2- Weak Axiom of Cost Minimization (WACM)

### Czech Republic

| Industry[a] | 1 | | 2 | | 3 | | 4 | | 5 | | 6 | | 7 | | 8 | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| **Year** | **WtNt [c]** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** |
| **2000-1999** | 1542651 | 1682887 | 1090976 | 1175858 | | | | | **1045279** | **1020248** | | | | | | |
| **1999-1998** | | | 1079690 | 1147171 | 948527 | 985969 | 721025 | 744739 | | | 2487936 | 2604558 | 1777254 | 1994616 | | |
| **1998-1997** | 1537031 | 1599166 | 1100584 | 1165324 | 908958 | 918490 | 690271 | 720628 | **923550** | **913226** | **2505474** | **2488950** | 1870924 | 1947469 | 577056 | 605442 |
| **1997-1996** | | | | | **818015** | **594920** | **642923** | **638307** | | | **2262200** | **1990736** | 1726403 | 1597727 | | |
| **1996-1995** | | | | | | | | | | | | | 1367672 | 1466332 | 399305 | 413153 |
| **1995-1994** | 830670 | 832852 | 803643 | 879901 | | | | | | | | | | | | |
| **1994-1993** | | | 768526 | 881243 | 412806 | 477068 | | | | | | | 1151548 | 1381624 | | |

### Hungary

| Industry[b] | 1 | | 2 | | 3 | | 4 | | 5 | | 6 | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| **Year** | **WtNt [c]** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** |
| **2001-2000** | | | | | 7575796 | 11050533 | **7552098** | **7284693** | | | **5784446** | **3994686** |
| **2000-1999** | | | 10335328 | 11489357 | 10074123 | 10535012 | | | | | | |
| **1999-1998** | **9449533** | **8721855** | | | **9123915** | **8765319** | **5543197** | **4347694** | | | | |
| **1998-1997** | | | | | | | | | 3693898 | 3779942 | | |
| **1997-1996** | 6651200 | 7032789 | | | | | 3944778 | 4029894 | | | | |
| **1996-1995** | 5890914 | 6252307 | 5720029 | 5959024 | 3271498 | 4134318 | 3262895 | 3509927 | 2447103 | 2500428 | | |
| **1995-1994** | 5211009 | 5649160 | | | | | | | | | | |
| **1994-1993** | 4720831 | 5152114 | | | 3802098 | 4078387 | | | 1794434 | 2090508 | | |

### Poland

| Industry[b] | 1 | | 2 | | 3 | | 4 | | 5 | | 6 | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| **Year** | **WtNt [c]** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** | **WtNt** | **WtNt-1** |
| **2000-1999** | | | | | | | | | | | | |
| **1999-1998** | | | | | 504772 | 529304 | 499214 | 543055 | 389383 | 436036 | 346662 | 375514 |
| **1998-1997** | **554018** | **538689** | | | | | | | 320777 | 337228 | | |
| **1997-1996** | | | | | | | | | | | 238314 | 244206 |
| **1996-1995** | **379264** | **281507** | **434540** | **331889** | **276341** | **229397** | **297214** | **238819** | | | | |
| **1995-1994** | | | | | | | | | | | 145655 | 147824 |

*Source: our calculations based on data from National Statistical Yearbooks (various years)*

[a] Czech Republic: 1=Food and tobacco; 2=Textiles, wearing apparel and leather; 3=Wood, paper and publishing; 4=Refined petroleum and chemicals; 5=Non-metallic products; 6=Basic metals and metal products; 7=Machinery and equipment; 8=Other manufacturing.

[b] Hungary and Poland: 1=Food products; 2=Total textile and wood activities; 3=Total petroleum, chemical, rubber, plastic products; 4=Total metal and mechanical products; 5=Total machinery, computers, RTV, communications; 6=Total vehicles and other transport equipments.

[c] $W_tN_t = w_t^s N_t^s + w_t^u N_t^u$; $W_tN_{t-1} = w_t^s N_{t-1}^s + w_t^u N_{t-1}^u$

**Table 3 – Fixed-Effect Estimation (Country*Sector Effects)**

| **Dependent Variable** | **WSh** | **WSh** |
|---|---|---|
| **Independent Variables** | **(1)** | **(2)** |
| $\ln(FDI)_{-1}$ | 0.001 | -0.001 |
| | (0.47) | (-0.85) |
| $\ln(X)_{-1}$ | -0.051*** | -0.022** |
| | (-4.75) | (-2.52) |
| $\ln(M)_{-1}$ | 0.002 | -0.024*** |
| | (0.20) | (-2.83) |
| $\ln(VA)_{-1}$ | 0.012 | -0.023* |
| | (0.67) | (-1.74) |
| $\ln(R\&D)_{-1}$ | 0.003 | 0.001 |
| | (0.76) | (0.31) |
| Year Dummies | YES*** | NO |
| Year*Country Dummies | NO | YES*** |
| F-Test for joint significance of Year Dummies (*p-value*) | 0.000 | - |
| F-Test for joint significance of Year*Country Dummies (*p-value*) | - | 0.000 |
| No. of obs. | 138 | 138 |
| F-Test for homogeneity of Country*Sector effects (*p-value*) | 0.000 | 0.000 |
| $R^2$ | 0.42 | 0.71 |

***t-statistics* in parenthesis. ***significant at 1%; ** significant at 5%; * significant at 10%**

**Table 4 – Fixed-Effect Instrumental Variable Estimation on the Pooled Sample (Country*Sector Effects)**

| Dependent Variables | WSh | | | $w^s / w^u$ |
|---|---|---|---|---|
| **Independent Variables** | **(1)** | **(2)** | **(3)** | **(4)** |
| $\ln(w^s / w^u)$PO_1 | 0.337$^{***}$ | 0.119 | 0.375$^{***}$ | - |
| | (4.43) | (1.04) | (10.01) | |
| $\ln(w^s / w^u)$CR_1 | 0.128 | 0.222 | 0.178$^{***}$ | - |
| | (1.24) | (1.31) | (3.04) | |
| $\ln(w^s / w^u)$HU_1 | 0.094 | 0.169 | 0.189$^{***}$ | - |
| | (0.65) | (1.31) | (3.22) | |
| ln(FDI)_1 | -0.005 | -0.004 | -0.005 | 0.034$^{***}$ |
| | (-1.18) | (-0.99) | (-1.09) | (3.98) |
| ln(X)_1 | -0.025$^{**}$ | -0.012 | -0.021$^{**}$ | -0.065$^{***}$ |
| | (-2.18) | (-1.08) | (-2.45) | (-2.84) |
| ln(M)_1 | 0.001 | -0.016 | 0.004 | -0.051$^{*}$ |
| | (0.14) | (-1.46) | (0.39) | (-1.95) |
| ln(VA) _1 | -0.021 | -0.034$^{**}$ | -0.026$^{**}$ | 0.103$^{***}$ |
| | (-1.41) | (-2.49) | (-2.11) | (3.02) |
| ln(R&D) _1 | -0.003 | -0.002 | -0.002 | 0.025$^{**}$ |
| | (-0.61) | (-0.49) | (-0.58) | (2.31) |
| Year Dummies | YES | NO | NO | YES$^{***}$ |
| Year*Country Dummies | NO | YES | NO | - |
| F-Test for joint significance of Year dummies (*p-value*) | 0.79 | - | - | 0.000 |
| F-Test for joint significance of Year*Country dummies (*p-value*) | - | 0.09 | - | - |
| Sargan's Test (*p-value)* | 0.72 | 0.21 | 0.80 | - |
| No. of obs. | 138 | 138 | 138 | 138 |
| F-Test for homogeneity of Country*Sector effects (*p-value*) | 0.000 | 0.000 | 0.000 | 0.000 |
| $R^2$ | 0.64 | 0.74 | 0.63 | 0.77 |
| $H_0$ : CRS *(p-value)* | 0.000 | 0.000 | 0.000 | - |
| WSh_Poland (fitted) | - | - | 0.502 | - |
| WSh_Czech Republic (fitted) | - | - | 0.316 | - |
| WSh_Hungary (fitted) | - | - | 0.297 | - |

***t-statistics* in parenthesis. $^{***}$significant at 1%; $^{**}$ significant at 5%; $^{*}$ significant at 10%**

**Table 5 – Fixed-Effect Instrumental Variable Estimation: Hungary and the Czech Republic (Country*Sector Effects)**

| Dependent Variables | WSh | | | $w^s / w^u$ |
|---|---|---|---|---|
| **Independent Variables** | **(1)** | **(2)** | **(3)** | **(4)** |
| $\ln(w^s / w^u)CR_{-1}$ | 0.214* | 0.235 | 0.170*** | - |
| | (1.77) | (1.31) | (2.94) | |
| $\ln(w^s / w^u)HU_{-1}$ | 0.241 | 0.148 | 0.186*** | - |
| | (1.43) | (1.13) | (3.19) | |
| $\ln(FDI)_{-1}$ | -0.004 | -0.003 | -0.003 | 0.017** |
| | (-1,04) | (-0.91) | (-0.82) | (2.17) |
| $\ln(X)_{-1}$ | -0.013 | -0.005 | -0.018* | -0.066** |
| | (-0.78) | (-0.37) | (-1.81) | (-2.35) |
| $\ln(M)_{-1}$ | -0.004 | -0.022* | -0.003 | -0.034 |
| | (-0.30) | (-1.66) | (-0.25) | (-1.08) |
| $\ln(VA)_{-1}$ | -0.031* | -0.042** | -0.025** | 0.075* |
| | (-1.68) | (-2.44) | (-1.91) | (1.95) |
| $\ln(R\&D)_{-1}$ | -0.003 | -0.001 | -0.001 | 0.021* |
| | (-0.63) | (-0.18) | (-0.32) | (1.88) |
| Year Dummies | YES | NO | NO | YES*** |
| Year*Country Dummies | NO | YES | NO | - |
| F-Test for joint significance of Year dummies (*p-value*) | 0.82 | - | - | 0.000 |
| F-Test for joint significance of Year*Country dummies (*p-value*) | - | 0.13 | - | - |
| Sargan's Test (*p-value*) | 0.57 | 0.13 | 0.74 | - |
| No. of obs. | 96 | 96 | 96 | 96 |
| F-Test for homogeneity of Country*Sector effects (*p-value*) | 0.000 | 0.000 | 0.000 | 0.000 |
| $R^2$ | 0.48 | 0.59 | 0.46 | 0.76 |
| $H_0$ : CRS (*p-value*) | 0.000 | 0.000 | 0.000 | - |
| WSh_Czech Republic (fitted) | - | - | 0.386 | - |
| WSh_Hungary (fitted) | - | - | 0.368 | - |

***t-statistics* in parenthesis. ***significant at 1%; ** significant at 5%; * significant at 10%**

**Table 6 – Fixed-Effect Instrumental Variable Estimation: Poland (Sector Effects)**

| **Dependent Variables** | **WSh** | | $w^s / w^u$ | **WSh ('95 and '96 excluded)** | $w^s / w^u$ **('95 and '96 excluded)** |
|---|---|---|---|---|---|
| **Independent Variables** | **(1)** | **(2)** | **(3)** | **(4)** | **(5)** |
| $\ln(w^s / w^u)$PO_1 | 0.246 | $0.306^{***}$ | - | 0.093 | - |
| | (1.15) | (5.03) | | (0.75) | |
| ln(FDI)_1 | 0.014 | 0.014 | $-0.073^{**}$ | 0.022 | -0.008 |
| | (0.56) | (1.30) | (-2.06) | (0.81) | (-0.10) |
| ln(X)_1 | -0.032 | -0.027 | 0.032 | $-0.037^{***}$ | -0.022 |
| | (-1.34) | (-1.47) | (0.82) | (-3.57) | (-0.73) |
| ln(M)_1 | -0.017 | -0.010 | 0.046 | 0.013 | 0.012 |
| | (-0.59) | (-0.49) | (0.84) | (0.88) | (0.27) |
| ln(VA) _1 | -0.028 | -0.034 | $0.281^{***}$ | $0.140^{***}$ | $0.242^{**}$ |
| | (-0.38) | (-0.73) | (2.71) | (3.64) | (2.18) |
| ln(R&D) _1 | -0.022 | -0.014 | -0.007 | $-0.030^{***}$ | -0.006 |
| | (-1.35) | (-0.99) | (-0.23) | (-3.72) | (-0.28) |
| Year Dummies | YES | NO | $\text{YES}^{***}$ | $\text{YES}^{***}$ | $\text{YES}^{*}$ |
| F-Test for joint significance of year dummies (*p-value*) | 0.41 | - | 0.000 | 0.000 | 0.06 |
| Sargan's Test (*p-value*) | 0.70 | 0.43 | 0.84 | 0.22 | 0.16 |
| No. of obs. | 42 | 42 | 42 | 30 | 30 |
| F-Test for homogeneity of sector effects (*p-value*) | 0.006 | 0.002 | 0.000 | 0.000 | 0.000 |
| $R^2$ | 0.85 | 0.80 | 0.93 | 0.97 | 0.95 |
| $H_0$ : CRS (*p-value*) | 0.33 | 0.22 | - | 0.08 | - |
| WSh_Poland (fitted) | - | 0.34 | - | 0.338 | - |

***t-statistics* in parenthesis. $^{***}$significant at 1%; $^{**}$ significant at 5%; $^{*}$ significant at 10%**