Foreign Direct Investment, Trade, and Skilled Labour Demand in Eastern Europe

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Abstract

In this paper, we study the effects of inward FDI and trade in final goods on relative skilled labour demand in Poland, Hungary and the Czech Republic. Our estimates show strong heterogeneity in the FDI effect across the three economies: the effect is positive, significant and sizeable in Poland; on the contrary, it is always insignificant in Hungary; finally, FDI exerts at most a very small negative effect in the Czech Republic. Such a heterogeneity is consistent with the different results emerging from existing theoretical models. Turning to trade, in each and every country, increasing exports of final goods lower relative skilled labour demand, as predicted by standard neoclassical trade theories.

JEL classification: F16, F23, J31

Keywords: Foreign Direct Investment; Trade; Wage inequality; Skilled labour demand

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1. Introduction

Since the fall of the Communist regime in 1989, Poland, Hungary and the Czech Republic have undertaken a rapid process of international integration, which has culminated in the EU membership in May 1, 2004. Although trade barriers and restrictions to inward Foreign Direct Investment (FDI) were already lifted before 1989, it is since then that trade flows with the EU Members and inward FDI from the same countries have risen sharply (Crinò, 2005). At the same time, the three countries have experienced sharp changes in their labour markets; in particular, wage inequality has worsened, and skilled labour has become progressively more rewarded. Are these two phenomena linked by some causal relationship? Many studies have already investigated this issue, focusing either on industrialised or on developing economies\(^1\). Yet, Poland, Hungary and the Czech Republic represent a peculiar case, because, due to the transition from the command to the market, they cannot be perfectly classified into either group.

Neoclassical trade theories and the standard Stolper-Samuelson theorem would suggest that increasing trade flows by unskill-abundant countries should lower relative demand for the skilled, thereby reducing wage inequality. Nevertheless, some of the existing experiences of trade liberalisation by developing economies show that wage inequality could actually rise with growing trade (Goldberg and Pavcnik, 2004).

The effect of FDI on relative skilled labour demand is more controversial. Generally speaking, two different sets of results emerge from the existing theoretical literature. According to the first, FDI can increase relative demand for skilled labour both at home and abroad. Feenstra

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and Hanson (1996b, 1997) develop a model in which FDI is described as a channel through which fragmentation of production takes place between countries with different relative skill endowments. Although the activities transferred by firms are unskill-intensive relative to the average task performed at home, they are more skill-intensive than the average task performed in the recipient country, and therefore, relative skilled labour demand rises in both economies as FDI takes place. The mechanism works because FDI shifts relative skilled labour demand outward; for a given supply, this increases both relative wage and relative employment of the skilled. We should notice, however, that such comparative static result may be weakened if the difference in skill endowments between investor and recipient country is not large. The effect of inward FDI may even reverse: if the unit cost curve in the North is steeper than in the South, increasing foreign capital in the South will push down the unit cost curve in this country, thereby inducing a decrease in relative demand for skilled labour (see Feenstra and Hanson (1997, fig.2, pp. 375)). According to the second set of results, the effects of FDI on relative skilled labour demand is even more ambiguous, both in the home and in the host economy, depending - among other things - on relative skill endowments. Markusen and Venables (1997) develop a general equilibrium model in which multinational firms (MNEs) operate in a monopolistically competitive sector, with a production process consisting of three activities: a firm-specific fixed cost using only skilled-labour, a plant-specific fixed-cost using both skilled and unskilled labour, and a branch-level final production stage which requires only unskilled labour. This framework gives rise to the following set of inequalities: activities carried out by the MNEs’ headquarters are more skill-intensive than those carried out by national firms, but the latter are in turn more skill-intensive than those carried out by the MNEs’ branches. Depending on the parameters of the model, as well as on the characteristics of the initial equilibrium and on the differences in
skill endowments between the two countries, FDI can either increase or decrease relative skilled labour demand abroad.

The lack of unambiguous theoretical implications on the role of FDI is reflected into the variety of conclusions reached by the existing empirical studies: while Feenstra and Hanson (1997) find clear evidence that U.S. FDI raised relative skilled labour demand in Mexico, Blonigen and Slaughter (2001) find no clear evidence of such an effect in the US\(^2\).

The sources of ambiguity in the effect of FDI on relative labour demand may be strong for our sample of countries, because they occupy some intermediate position in terms of relative skill abundance between developed and developing countries. Hence, there is clear need of an empirical analysis assessing the effect of FDI and trade on relative skilled labour demand and wage inequality in Poland, Hungary and the Czech Republic. Up to now, empirical evidence on these countries is limited. To the best of our knowledge, only two studies (Skuratowicz, 2001; Lorentowicz et al., 2005) have analysed the effect of FDI, focusing only on Poland: results show a positive and significant effect on relative skilled labour demand\(^3\). Another study by Egger and Stehrer (2003) has considered the effect of outsourcing: results show that outsourcing reduced relative labour demand for the skilled in all the three countries.

This paper tries to improve upon existing literature, by analyzing for the first time the effects of inward FDI jointly on the three economies, and by comparing these effects with those of trade in final goods. To this purpose, we use an industry/year panel covering 6 manufacturing

\(^2\)A related stream of literature focuses instead on the effect of outward FDI on domestic relative skilled labour demand. Also in this case, results are ambiguous: while Head and Ries (2002) and Hansson (2005) find positive effects of foreign affiliates activities on relative skilled labour demand in Japan and Sweden, Slaughter (2000) finds no such an effects for the U.S.. See Crinò (2006a) for a recent survey.

\(^3\)Bedi and Cieslik (2002) ask a different question: whether increasing inward FDI raised average wages and average wage growth in Poland. They find that this was indeed the case, as workers in industries with greater foreign presence enjoyed both higher wages and higher wage growth.
sectors over the period 1994-2002, and estimate a skilled labour share-equation derived from a short-run translog cost function; while doing so, we also tackle some important estimation issues concerning the treatment of endogenous and predetermined regressors in share equations with latent heterogeneity, thereby providing a second improvement upon existing literature.

To preview the results, we find significant heterogeneity across the three countries as far as the effects of FDI are concerned: in particular, the effect is significant and positive in Poland; in the Czech Republic, it is generally negative, although very small in size and not always robust across specifications; finally, in Hungary, the effect is always insignificant. As far as trade is concerned, results are instead much more homogeneous; in all countries, we find negative and significant effects of exports on the relative skilled labour demand: a one percent increase in exports causes a loss between 3 and 4 percentage points in the skilled labour share.

We justify the heterogeneity in the FDI effect in the light of the relative positions of the three countries with respect to the main investors (EU Members) in terms of skill endowments: Poland is more unskill-abundant than Hungary and the Czech Republic relative to the EU Members (European Commission, 2004), and therefore it is in Poland that FDI may have played a stronger role. The result for trade is instead consistent with the standard predictions from neoclassical trade theory.

The remainder of the paper is organised as follows: in section 2 we present the data and some stylised facts; section 3 develops the empirical model and section 4 presents the results. Section 5 finally concludes.
2. Data description and stylised facts

Highly disaggregated and comparable industry-level data on inward FDI in Poland, Hungary and the Czech Republic are not available for a sufficiently long time span. To our knowledge, the best available compromise between the cross-section and the time-series dimension of FDI data for these countries is the OECD “International Direct Investment Statistics”, which allows to focus on 6 manufacturing industries over the period 1994-2002. From this dataset, we choose the inward FDI position (stock) as a proxy of MNEs penetration. For the Czech Republic, however, OECD data on inward stocks are available only for a few years; we therefore use data on FDI inflows provided by the “Balance of Payments Statistics” of the Czech National Bank, and retrieve FDI stocks as cumulative inflows. Data on employment and wages for skilled and unskilled workers come from each country’s statistical yearbooks; we follow the usual skill approximation and accordingly define non-manual and manual employees as skilled and unskilled workers. Data on exports and imports of final goods come from the OECD "Stan Database for Industrial Analysis". Finally, in order to take into account the effect of technical progress in explaining the rise in wage inequality (Berman et al., 1994), our database includes also the business enterprise expenditure on R&D, also retrieved from the OECD Stan Database. All data are converted in constant 2000 prices, using the GDP deflator provided by the OECD "Economic Outlook Database". Descriptive statistics on these variables are reported in Table 1; the appendix provides a more detailed description of the dataset.

Some stylised facts on these data are as follows. Poland, Hungary and the Czech Republic have experienced rapid increases in wage inequality during the 1990s: between 1994 and 2002,

\footnote{We are indebted to R. Stehrer for the provision of the dataset up to 1999.}
the relative wage of skilled workers has increased from 1.47 to 1.82 in the Czech Republic, from 1.88 to 2.09 in Hungary and from 1.40 to 1.92 in Poland. At the same time, however, relative employment of skilled workers – as proxied by the ratio between the number of non-manual and manual employees - has declined in Hungary and in the Czech Republic (from 0.32 to 0.29 and from 0.42 to 0.36, respectively); in Poland, instead, it has risen from 0.3 to 0.33 (Figure 1). The skilled labour shares of total manufacturing wage bill and employment have moved consistently with these patterns: the skilled labour share of total wage bill has risen from 0.37 to 0.372 in Hungary, from 0.29 to 0.38 in Poland and from 0.37 to 0.39 in the Czech Republic; the skilled labour share of total employment, instead, has increased only in Poland (from 0.23 to 0.25), and decreased in the other two countries (from 0.24 to 0.22 in Hungary and from 0.29 to 0.26 in the Czech Republic)\(^5\). Decomposing the changes in the skilled labour shares of total employment and wage bill, Crinò (2005) finds that they are mostly due to within-industry rather than to between-industries variations\(^6\). Therefore, whatever the causes of wage inequality, they seem to act mostly 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 developed and developing countries (Berman \textit{et al.}, 1998)\(^7\). Nevertheless, Crinò (2005) also shows another interesting piece of evidence: between-industry variations in the skilled labour share of total

\(^5\)Looking at the dynamics of relative wage and employment at the 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, 2005).

\(^6\)The decomposition follows Berman \textit{et al.} (1994).

\(^7\)Note that Crinò’s results are obtained on the whole set of 14 NACE Rev.2 manufacturing industries. The sample we use here, instead, excludes the sectors "other non-metallic mineral products" (DI), "electrical and optical equipments" (DL) and "manufacturing not elsewhere classified" (DN) due to the need of obtaining wage and employment estimates with the same breakdown as the FDI data.
employment have been negative in all countries; similarly, between-industry variations in the skilled labour share of total wage bill have been negative in Hungary and in Poland. This evidence suggests that some forces - in particular, trade in final goods - have indeed acted in favor of the unskilled, by increasing the weight of unskilled labour-intensive industries on total manufacturing employment.

Summing up, wage inequality has worsened in each and every country. However, the channels through which this has happened differ. In Poland, both relative wages and relative employment of the skilled have risen. Hence, Poland seems to have experienced an outward shift in relative skilled labour demand, which has mostly occurred within-industry. This evidence leaves room for a positive effect of FDI; at the same time, the between / within-industry decomposition supports also a role for trade. On the contrary, evidence for Hungary and the Czech Republic is at odds with an outward shift in relative skilled labour demand, since relative wages have grown, but relative employment has declined; however, also for these countries there is evidence supporting the role of between-industry demand shifters that have acted in favor of the unskilled.

Over the same period, Poland, Hungary and the Czech Republic have received the largest FDI inflow among the CEECs; along with Slovakia and the Russian Federation, they account today for 3/4 of the region’s inflows. In percentage of GDP, economy-wide FDI stocks have sharply increased in the last decade, rising from values below 4% in 1990 to more than 40% in Hungary and in the Czech Republic and to more than 20% in Poland in 2000 (UNCTAD, 2002). Consistently, FDI stocks have markedly increased also in the manufacturing sector, which accounts for nearly half of the overall 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, 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 and transport equipment; in Hungary and Poland, the food industry still accounts for a share around 30%.

FDI is just one expression of the rapid process of international integration which the three 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 (Crinò, 2005). Moreover, 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 labour, and on the other hand, of the improvements obtained in intra-branch product quality (Landesmann and Stehrer, 2002).

In conclusion, since the beginning of the transition, 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 labour has become more rewarded. The contemporaneous occurrence of these phenomena, however, does not necessarily imply a causal relationship between them. The remaining part of the paper, therefore, investigates the existence of such a

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8 For a more detailed industry-level analysis, see Crinò (2005).
9 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.
causal link, by making use of various methods of multivariate regression.

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 minimisation within sectors for given stock of capital. In more detail, we assume the existence of a representative firm for a given industry, which minimises the cost of skilled and unskilled labour to produce a given amount of output, treating capital as fixed over the relevant sample period. This optimisation problem for the firm yields the following variable cost function

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

(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.

The baseline parametric model is derived by the following translog function for (3.1):

\[
\ln(VC) = \beta_0 + \sum_{h=s,u} \beta_h \ln w^h + \beta_Y \ln Y + \beta_K \ln K
+ \frac{1}{2} \left( \sum_{h=s,u} \beta_{hh} \ln^2 w^h + \beta_{yy} \ln^2 Y + \beta_{kk} \ln^2 K \right) +
+ \beta_{su} \ln w^s \ln w^u + \sum_{h=s,u} \beta_{yh} \ln w^h \ln Y + \sum_{h=s,u} \beta_{kh} \ln w^h \ln K
\]

(3.2)

One of the advantages of starting from a functional form for \( C(w^s, w^u, Y, K) \), rather than
is the possibility of identifying input price elasticities from estimated input demand systems or share equations with a minimum amount of computational difficulties. By applying Shephard Lemma and exploiting the homogeneity and adding-up restrictions for the translog, we can derive the following share equation for skilled labour:

\[
W_{Shs} = \beta_s + \beta_{ss} \ln(w) + \beta_{Ks} \ln K + \beta_{Ys} \ln Y
\]  

(3.3)

where \(W_{Shs}\) is the skilled-labour share of total industry wage-bill and \(w\) denotes \(w^s/w^u\). Modifications of equation (3.3) will be estimated and tested using a battery of estimators on the pooled data for the three countries.

4. Results

For estimation\(^{10}\) equation (3.3) is expanded to accommodate the effects of several variables on relative skilled labour demand. First of all, we use inward-FDI stocks \((FDI)\) as a measure of foreign penetration. Moreover, we account for the influence of international trade by using exports and imports of final goods \((X\) and \(M\), respectively). We also control for technological progress by means of the total business enterprise expenditure on R&D \((R&D)\). Finally, industry output, \(Y\), is measured by real production. Since there are no observations available on the capital stock at the industry level, we assume that all variations of capital stock correlated with the included explanatory variables are captured by both country-sector and time specific effects. This gives us the following baseline estimating share equation:

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\(^{10}\)Details on the computing codes used in the paper are relegated into the computational appendix at the end.
$$WSh_{i,j,t} = \beta_w \ln(w_{i,j,t}) + \beta_{1,j} \ln(FDI_{i,j,t})$$
$$+ \beta_2 \ln(X_{i,j,t}) + \beta_3 \ln(M_{i,j,t}) + \beta_4 \ln(Y_{i,j,t})$$
$$+ \beta_5 \ln(R&D_{i,j,t}) + \alpha_{i,j} + \gamma_t + \varepsilon_{i,j,t}$$

(4.1)

where $\varepsilon_{i,j,t}$ is an idiosyncratic error term, $\alpha_{i,j}$ is a country-sector specific time-invariant effect and $\gamma_t$ is a time-specific effect that is invariant across country-sectors groups. $i = 1, ..., 6$ indexes the manufacturing industries, $j \in \{hu, po, rc\}$ indexes the countries, and $t = 1994, ... 2002$ is the time span of our panel.\textsuperscript{11}

Including both the cross-sectional effect $\alpha_{i,j}$ and the time specific effect $\gamma_t$ into the share equation is beneficial for two main reasons. The first is well-known and has to do with the removing of unobserved heterogeneity from the systematic part of the specification. The $\alpha_{i,j}$ term accommodates permanent latent shocks varying across countries and sectors that may be arbitrarily related to the observed regressors. Such extension is especially important for the countries examined in this paper, where substantial differences are found regarding both the period of beginning and the average intensity of the transition process. Another piece of unobserved heterogeneity is captured by the $\gamma_t$ term. This comprises all transitory latent shocks at the aggregate level that hit the labour demand and are also arbitrarily related to trade and FDI flows. There are many factors that fall in this category, ranging from the degree of trade integration between the three countries to more aggregate effects, such as world trade agreements.

\textsuperscript{11}In order to pool the data meaningfully, we convert all variables in US$, using the PPP exchange rates provided by the OECD "Structural Statistics for Industry and Services".

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and all available achievements of technical progress.

The second benefit from including fixed effects, indeed not so often emphasised in the applied literature but still important, has to do with the non-systematic part of the regression model. In fact, least squares estimators when applied to equation (4.1), while efficiently purging the systematic part of the regression from the unobserved heterogeneity terms $\alpha_{i,j}$ and $\gamma_t$, remove also all group-specific and time-specific effects in the composite error ($\varepsilon_{i,j,t} + \alpha_{i,j} + \gamma_t$), so that inference is robust to *constant* patterns of cross-sectional and serial correlation in the covariance matrix of the composite error (Wooldridge, 2002). This, however, does not ensure against *arbitrary* serial correlation in the realisations of $\varepsilon_{i,j,t}$ and, consequently, against the bias in the standard error estimates that would arise. Therefore, it is crucial for an adequate inference strategy 1) that the presence of arbitrary patterns of serial correlation in estimating equations be carefully detected and 2) if serial correlation is found, that a consistent estimator for the variance-covariance matrix of the idiosyncratic term be applied.

We follow two simple approaches to testing serial correlation in panel data: the Wooldridge’s (2002) test of first-order serial correlation and the Arellano and Bond (1991) (AB, henceforth) approach for testing serial correlation of any order. The former is simpler but it is applicable only in the presence of strictly exogenous regressors; besides, it is designed for the specific null hypothesis of zero first-order autocorrelation. The latter, instead, is implementable in the presence of endogenous regressors and allows to perform separate tests for any order of serial correlation until the maximum allowed by the time series dimension of the data. Both tests are carried through first-difference residuals\(^{12}\). We have applied the two tests to various

\[^{12}\text{Under the null hypothesis of zero serial correlation in the errors in levels, there is first order serial correlation in first-differenced errors and this equals -0.5. So, the Wooldridge statistics tests the null by checking that first-order autocorrelation in first-difference residuals be not significantly different from -0.5. Also, under the null of} \]
specifications of equation (4.1) and for different estimators (as presented below); results are reported in Table 2. While both tests confirm the presence of first-order serial correlation in the idiosyncratic error (AB at 1% and Wooldridge at any conventional significance level), the AB test does not support higher than first order autocorrelation. The former conclusion, however, is enough to motivate use of a robust estimator for the variance-covariance matrix. So, for all specifications and estimators of this paper standard error estimates have been corrected through the White’s procedure, in the version of the clustered covariance estimator suggested by Arellano (1987). The cluster estimator in our case produces covariance estimates that are robust to arbitrary patterns of serial correlation within country-sector groups as well as groupwise heteroskedasticity\textsuperscript{13}.

An F-test always rejected the restriction of equal impact of FDI on the skilled labour share of total wage bill across the three countries (see Table 3 below); this is in line with the predictions of existing theoretical models. Hence, in all our estimating equations, we leave the coefficient on FDI vary across countries\textsuperscript{14}.

4.1. Fixed effect estimation

Several studies on the skill upgrading (Machin and Van Reenen, 1998; Blonigen and Slaughter, 2001; Pavcnik, 2003; Lorentowicz et al., 2005) have noted that the log-relative wage is likely to be endogenous in a share equation. In fact, although part of the variation in relative wages zero serial correlation of order \( p \) in the errors in levels, we may expect serial correlation of order \( p + 1 \) in first-differenced errors but not \( p + 2 \). So, the appropriate AB statistics for that null detects that the \( (p + 2)^{th} \)-order autocorrelation coefficient, estimated from first-difference residuals, be not significantly different from zero.

\textsuperscript{13}In a recent paper Stock and Watson (2006) demonstrate that the standard White correction to heteroskedasticity in fixed effect panel data models is inconsistent for fixed \( T \). They also show that the inconsistency does not carry over into the clustered covariance estimator used in this paper.

\textsuperscript{14}We show below also results from regressions with country-specific coefficients on exports. As it will emerge, the restriction of equal exports coefficients is accepted by the data.
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 in a reduced-form version of (4.1). Other studies have also extended such a reduced-form with additional variables, controlling for exogenous shifts to relative labour supply, which may impact the skilled labour share of wage bill: for example, Feenstra and Hanson (1997) include the log alternative wage for both types of workers among the regressors. Following this approach, in our first specification (MODEL 0a) we omit the relative wage variable and estimate equation (4.1) by a Within estimator allowing for unobserved heterogeneity across industries; moreover, we also estimate equation (4.1) by including a proxy for the log alternative relative wage of the skilled, $\ln \left( w^o \right)_{j,t}$ (MODEL 0b)\(^{15}\). Estimated parameters are presented in Table 3 (column 1 and 2). The total number of observations is 162\(^{16}\). Results from the two specifications do not differ qualitatively from each other, as expected in the light of the lack of significance of the log alternative relative wage in column 2. The fixed time effects (unreported) are always highly significant. Estimated coefficient on FDI is positive and highly significant for Poland, positive but only slightly significant for Hungary, negative and insignificant for the Czech Republic. Turning to the trade coefficients, estimated parameters are negative for both exports and imports, and also significant for exports. Finally, the coefficient on output is insignificant, revealing the presence of constant returns to scale in production; no significant effect emerges for R&D, suggesting that technological change has been skill-neutral. These results are in line with

\(^{15}\) We follow Feenstra and Hanson (1997) in defining the alternative wage for the skilled as the wage paid in the tertiary sector and the relative wage for the unskilled as the wage paid in the primary sector. Notice that the alternative relative wage does not vary across industries.

\(^{16}\) Our panel consists in fact of 6 manufacturing industries observed for 9 years in each country.
our expectations: the effect of FDI is heterogeneous across countries, consistently with existing theoretical and empirical results and with our stylised facts; trade in final goods, instead, clearly reduces relative demand for skilled labour, in line with neoclassical trade theory and with our stylised facts.

The estimates presented above must be interpreted with care. In the estimation of a reduced-form equation as the one above, coefficient estimates will capture both demand-side and supply-side effects, thereby obscuring the demand-side impact of FDI and trade flows. A structural interpretation of the estimated elasticities from a share equation excluding the relative wage could be justified only under the assumption that the portion of the relative wage correlated with the explanatory variables can be fully captured by the inclusion of individual and time dummies. In our case, however, this assumption seems quite unreasonable: indeed, Crinò (2005, Fig. 4, pp. 32) shows sizeable cross-industry variations in relative wages, which are unlikely to be captured by uncorrelated idiosyncratic sectoral shocks. For this reason, we choose to base our following analysis on estimated parameters from the structural demand equation in (4.1), which includes also the log-relative wage among the regressors (MODEL 1, column 3). As expected, the coefficient on \( \ln (w)_{i,j,t} \) turns out to be highly significant.

The significantly positive impact of the log-relative wage on the skill labour share may cause some concern about the regularity of the underlying demand equations for skilled and unskilled workers. Indeed, it is well known that regularity conditions for share equations derived from the translog cost function are only met locally. So, what really matters for our empirical analysis is that the estimated coefficient on \( \ln (w)_{i,j,t} \) be not so high as to reject a negatively sloped demand curve over a relevant portion of the sample. In this view, a natural requirement is that the estimated demand curves be negatively sloped over a neighborhood of the sample averages. This
occurs if and only if the Hessian of the underlying translog cost function is negative semidefinite at the average sample point or equivalently, in our two-inputs case, if and only if the matrix of the substitution elasticities is negative semidefinite at the average share $W_{Sh}$, that is for all $\beta_w$ such that $\epsilon_{s,u} = \frac{\beta_w + W_{Sh}}{W_{Sh}} \leq 0$, with $W_{Sh} = 0.365$\(^{17}\). Hence, the estimated skill labour demand curve will be negatively sloped over a neighborhood of $W_{Sh} = 0.365$ for all $\beta_w \leq 0.231$.

Evidently, the estimated relative wage coefficient is below this threshold. This can also be tested by means of a one-tailed t-test for the null hypothesis that $\beta_w \leq 0.231$; results from this test are reported in Tables 3 and 4 (below) for all models and estimators, confirming that our estimated parameters are always consistent with a locally well-behaved cost function. The inclusion of the relative wage produces some noticeable changes in the coefficients of interest: namely, $FDI$ now loses significance in Hungary and gains significance in the Czech Republic, where the absolute size of the coefficient remains however very low; $FDI$ remains instead significant and positive in Poland, although the size of the coefficient almost halves. Trade in final goods continues to produce the effects predicted by standard neoclassical theories: the coefficient on exports, in fact, is again significant and negative, although its absolute size shrinks. Hence, reduced-form estimation may lead to overstate the effects of the variables of interest. Finally, no significant changes occur in the remaining coefficients.

In the last column of Table 3, we let the coefficients on exports vary across countries (MODEL 2). If the trade patterns of the three economies were different, in fact, restricting the coefficients to be equal across countries would not be realistic. Moreover, since exports are typically correlated with FDI, gauging the effect of foreign penetration in each economy requires to allow for

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\(^{17}\)This value is obtained by averaging across the three countries the mean values for the skilled labor share of total wage bill reported in Table 1.
country specificity also in the export variable. Our theoretical prior, however, is that the effect of exports should not differ significantly across these three economies: all of them, indeed, are relatively less endowed with skilled labour than the EU Members (the main trading partners) and have witnessed similar, though by no means identical, changes in their export patterns since the fall of the Communist regime (Crinò, 2005). And, in fact, the F-test for the null hypothesis of equality of the export coefficients reported in the third column of Table 3 supports our expectation. As a consequence, results on $FDI$ remain robust to the inclusion of country-specific exports, in terms of both sign, significance and absolute size of the coefficients. Coefficients on exports are significantly negative for all the three countries and show similar size. Finally, remaining explanatory variables are again insignificant.

Summing up, our results have insofar shown that the effects of FDI on relative labour demand are heterogeneous across the three countries. More specifically, the only case in which increasing foreign penetration has been found to raise relative skilled labour demand is Poland; in the other two countries, $FDI$ either is insignificant (Hungary) or exerts very limited negative effects (Czech Republic). Such heterogeneity of results reflects the conflicting predictions from existing theoretical models and is consistent with the inconclusive evidence emerging from previous empirical studies. The question is: why does FDI exert positive effects in Poland and not in the other two countries? One possible explanation can be found by looking at the position of Poland in terms of skill endowment relative to Hungary and the Czech Republic: while all countries are definitely more unskill-abundant than the EU Members, Poland is relatively more so as compared to Hungary and the Czech Republic. According to the European Commission (2004, Table 42, p. 111), in 2003, only 20% of non-manual workers were employed in Poland, accounting for less than 39% of total employment. These figures were higher in Hungary and in the Czech
Republic in the same year: in Hungary, the employment rate of non-manual workers reached 24% and non-manual employees accounted for 42% of total employment; in the Czech Republic, the employment rate of non-manual workers reached 29% and these employees accounted for nearly 45% of total employment. For this reason, FDI is more likely to exert a positive effect on relative skilled labour demand in Poland than in Hungary and the Czech Republic: according to existing theoretical results, in fact, in order for FDI to shift relative skilled labour demand outward in the recipient country, the latter has to be significantly more endowed with unskilled labour than the investing economy; otherwise, the effect may weaken or even reverse. This is confirmed by previous empirical studies: while Feenstra and Hanson (1997) find strong evidence of positive effects of U.S. FDI in Mexico, Blonigen and Slaughter (2001) find that Japanese FDI in the U.S. either leaves unaffected or lowers relative skilled labour demand. Hence, the more countries differ in skill endowments, the stronger is the effect of FDI. Other studies exist that confirm our results on Poland: Skuratowicz (2001) and Lorentowicz, et al. (2005), in fact, find that inward FDI significantly raised relative skilled labour demand and wage inequality in this country during the 1990s.

Turning to trade in final goods, our results show that it acts as predicted by standard neoclassical theories: Poland, Hungary and the Czech Republic, relatively more abundant of unskilled labour as compared with their main trading partners, specialise in unskill-intensive productions, and this tends to lower relative demand for skilled labour and wage inequality. Estimated coefficients imply that a one percent increase in exports causes a loss between 3 and 4 percentage points in the skilled labour share. Also in this case, results are supported by our stylised facts and by previous findings on the same economies (Egger and Stehrer, 2003).
4.2. Robustness Analysis: IV Estimation

We now move to investigate the robustness of our results. The above estimates, although satisfactory in many respects, fail to address the possible endogeneity of the relative wage at the country-sector level. If this were the case, estimated coefficients would be inconsistent. Hence, in order to check the robustness of our results to the endogeneity of $\ln(w)_{i,j,t}$, we reestimate both MODEL 1 and MODEL 2 using a battery of Instrumental Variables (IV) estimators.

Instrumental variables have been selected and tested according to the two criteria of *instrument relevance* and *instrument validity*. Instrument relevance requires that each instrument be highly correlated to the endogenous regressors. When this is not the case the instrument is said to be weak and the IV estimates will be biased in the same direction of the OLS estimates (see Staiger and Stock, 1997). Importantly, instrument relevance can be tested empirically. A simple test, suggested by Bound *et al.* (1995), is implementable in the presence of one endogenous regressor. It looks at the squared correlation between the included endogenous and its prediction from the first-stage equation after partialling out all the included exogenous regressors. It is implemented as an F test of joint significance of the instruments in the first-stage regression. Staiger and Stock (1997) note, though, that the problem of weak instruments can arise even when the null of zero partial correlation can be rejected at conventional levels of significance (5% or even 1%) and regardless of the sample size. As a rule of thumb, they suggest that a value of the first-stage F test less than 10 should raise concern about the relevance of the chosen instruments. In this paper we abide by this simple rule to assess instrument relevance$^{18}$. The

\footnote{Stock and Yogo (2002) stated the null hypothesis of weak instruments in rigorous terms tabulating the critical values for the F-test, so providing solid theoretical foundations to the Staiger and Stock’s rule of thumb. Their inference strategy, however, is designed under the assumption of a spherical covariance matrix of the idiosyncratic error, so it is not suitable for our empirical analysis.}
second criterion is that of instrument validity. It requires the absence of correlation between each component of the instrument set and the idiosyncratic error. While the restrictions placed by instrument validity for the exact identification of the model are not testable, the overidentifying restrictions it implies when the number of instruments exceeds that of endogenous regressors can be tested by means of the Sargan statistic, or its robust alternative given by the Hansen J statistic. In order to ease the evaluation of our estimation results, for each model and estimator we state explicitly the assumptions that guarantee instrument validity and whether the model is just identified or overidentified. In the latter case, then, we report in the tables the probability values of the Hansen J statistics.

IV results are reported in Table 4. In each column we show the first-stage F tests. All our proposed IV estimates show high values for these statistics, well beyond the rule of thumb of 10. We also report the F-test for equality of FDI coefficients across countries: as before, such a restriction is not supported by the data, thereby validating the use of country-specific FDI slopes. Finally, local regularity conditions for cost minimisation are satisfied in each and every specification, as the one-tailed test for the null \( \beta_w \leq 0.231 \) suggests.

The first four columns of Table 4 report estimated parameters from MODEL 1, that is, the specification with cross-country equality of the exports coefficients imposed; the remaining part of the table reports parameter estimates obtained by relaxing this restriction (MODEL 2).

4.2.1. TSLS on Within transformations

In column 1 of Table 4, we present results for our first IV estimator. It is a Two Stage Least Square (TSLS) estimator where latent heterogeneity across country-sectors is removed by the within transformation: given any variable \( z_{i,j,t} \) its within transform is \( z_{i,j,t}^* = z_{i,j,t} - \bar{z}_{i,j} \) where
\[ z_{i,j} = \frac{1}{T} \sum_{t=1}^{T} z_{i,j,t}. \] Throughout, we will refer to this estimator as the Within Two Stage Least Square estimator (WTSLS).

Let \( Z \) denote the vector of all included regressors other than endogenous ones:

\[
Z_{i,j,t} = \left[ \ln(FDI)_{i,j,t}, \ln(X)_{i,j,t}, \ln(M)_{i,j,t}, \ln(Y)_{i,j,t}, \ln(R&D)_{i,j,t} \right].
\]

and \( Z_{i,j} = [Z_{i,j,1}, Z_{i,j,2}, \ldots, Z_{i,j,T}] \). Consistency of WTSLS requires that all the components of \( Z \) be strictly exogenous conditional on \( \alpha_{i,j} \) and \( \gamma_t \): \( E(\varepsilon_{i,j,t}|Z_{i,j,\alpha_{i,j},\gamma_t}) = 0 \). While conditional strictly exogeneity rules out any feedback from the idiosyncratic shock to \( Z_{i,j} \), it leaves unrestricted the correlation between \( Z_{i,j} \), \( \alpha_{i,j} \) and \( \gamma_t \); so that \( Z_{i,j} \) is allowed to respond to 1) transitory aggregate labour demand shocks (the time effects \( \gamma_t \)); and 2) permanent sector-specific labour demand shocks (the group effects \( \alpha_{i,j} \)). Under conditional strict exogeneity, all usable lags of the within transforms \( Z_{i,j}^* \) are valid instrument for \( \ln(w)_{i,j,t} \). 19.

After applying WTSLS to different instrument sets we have not been able to find any relevant difference in results across the various versions of WTSLS as well as in comparison to the Within estimators. We report and comment here only results peculiar to an exactly identified WTLSL estimator using \( \ln(Y)_{i,j,t-2} \) as instrument20. The F-test for instruments relevance is equal to 17.6, which can be taken as evidence that our instrument is strong. Results prove robust to the use of WTSLS: if compared with estimates in column 3 of Table 3, estimated coefficients keep their sign and significance: \( FDI \) is significantly positive in Poland, significantly negative in the Czech Republic (although the absolute size of the effect is very small) and insignificant.

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19 In order not to lose additional observations, for all instruments used in this paper we set to zero the missing values generated by the lag operator.

20 All unreported estimates mentioned in the paper are available from the authors on request.
in Hungary; exports are significantly negative, whereas the remaining explanatory variables are
insignificant. Estimated parameters appear robust also in terms of their absolute size: with
the only relevant exception of the relative wage, whose coefficient increases sizably relative to
column 3 in Table 3, the other parameters have almost the same size as before.

The high first-stage F-test values, along with our parsimonious choice of moment conditions,
makes us expect that, conditional on instrument validity, the finite-sample bias in our estimates
should not be much of a problem, notwithstanding the small number of cross-sectional units
available. In order to gain more insight into this issue we have used results in Bruno (2006)
to estimate the approximation of the finite sample bias for an overidentified WTSLS using
\( \ln(M)_{i,j,t-2}^* \) and \( \ln(X)_{i,j,t-2}^* \) as instruments, in addition to \( \ln(Y)_{i,j,t-2}^* \). Coefficient estimates
and test statistics are basically the same as in the just identified WTSLS; the Hansen J statistics
does not reject the implied overidentifying restrictions; more to the point, the estimated bias
approximations turn out to be virtually zero, so confirming our expectation.

4.2.2. TSLS on Forward-Orthogonal-Deviations

As is well known (see Ziliak, 1997 among others), for WTSLS to be consistent the choice of the
instruments must be restricted to variables that are strictly exogenous to the error term. In

\[ \text{Bruno (2006) extends the Nagar’s bias approximations derived in Buse (1992) to the overidentified WTSLS with AR(1) idiosyncratic errors. Monte Carlo experiments therein shows that bias approximation estimates are often an accurate measure of the actual bias. Bias approximations have been originally employed in the time series literature and then introduced into the dynamic panel data literature by Kiviet (1995) and further developed, among others, by Bruno (2005a,b).} \]

\[ \text{To conserve space these results are not presented here, but are available from the authors on request.} \]

\[ \text{To illustrate, consider a variable } Z \text{ that is only predetermined to the error term, that is } \mathbb{E}(\varepsilon_{i,j,t}|Z_{i,j,1},...,Z_{i,j,t},\alpha_{i,j},\gamma_{i}) = 0, \text{ then WTSLS with current or past values of } Z \text{ as instruments would not offer consistent estimators of the coefficients, since the within transformation would bring all past realizations of the error back into the current realization of its within transform, } \varepsilon_{i,j,t}^* \equiv \varepsilon_{i,j,t} - \tau_{i,j}; \text{ via the error group mean } \tau_{i,j} \equiv \frac{1}{T} \sum_{t=1}^{T} \varepsilon_{i,j,t}; \text{ and this whether or not the values of } Z \text{ are within-transformed.} \]
this respect, the forward orthogonal transformation suggested by Arellano and Bover (1995)\(^{24}\) 
is more convenient than the within transformation for removing the latent heterogeneity components, 
since valid instruments are obtained also as lagged values of predetermined, or even endogenous, regressors. The TSLS estimator applied to the model transformed in forward orthogonal deviations (FODTSLS) has been discussed and evaluated by Ziliak (1997). Besides its flexible treatment of instruments, it also recommends itself for its satisfactory finite sample properties compared to GMM estimators. For these reasons we supplement our analysis with various applications of FODTSLS.

For instrument validity we maintain the following sequential moment condition:

\[
E\left( \varepsilon_{i,j,t} | \ln (w)_{i,j,1}, ..., \ln (w)_{i,j,t-p}, Z_{i,j}, \alpha_{i,j}, \gamma_t \right) = 0
\]

(4.2)

for all \(t = p + 1, ..., T\) and some fixed \(0 \leq p \leq T - 1\).

The formulation of condition (4.2) is general enough to leave open a broad range of possibilities for the underlying Data Generation Process (DGP) of the relative wage. For \(p \geq 1\) not only does it permit feedbacks from all lags of \(\varepsilon_{i,j}\) to \(\ln (w)_{i,j,t}\) and a contemporaneous feedback between \(\varepsilon_{i,j,t}\) and \(\ln (w)_{i,j,t}\), but also feedbacks in the opposite direction from lags more recent than \(\ln (w)_{i,j,t-p}\) to \(\varepsilon_{i,j,t}\). Indeed, either in its most restrictive form of \(p = 0\) (\(\ln (w)_{i,j,t}\) is predetermined) or in the weaker form of \(p \geq 1\) (\(\ln (w)_{i,j,t}\) is endogenous and possibly "post-determined") condition (4.2) governs, although often implicitly, the treatment of the endogenous regressors. \(24\) The forward orthogonal transformation takes all variables in the model in weighted deviations from the forward group means: given \(x_{i,j,t}\) it has \(\tilde{x}_{i,j,t} = c_t (x_{i,j,t} - \bar{x}_{i,j,t})\) where \(\bar{x}_{i,j,t} \equiv \frac{1}{T} \sum_{s=t+1}^{T} x_{i,j,s}\) and \(c_t^2 = (T - t) / (T - t + 1)\). The presence of the weight \(c_t\) ensures that if the idiosyncratic error is homoskedastic and not serially correlated then also its FOD transform is homoskedastic and not serially correlated (see Arellano, 2003).
right-hand variables in most recent panel data studies based on IV estimators (Ziliak, 1997, for example, focuses only on the case of $p = 0$; for a more general treatment see Arellano, 2003).

But the question arises: How to choose the value for $p$? Put it differently, is there any empirical justification for limiting the degree of "post-determination" of the relative wage to a given value? Our solution is along the following lines. As long as the DGP for $\ln (w)_{i,j,1}$ is unknown, condition (4.2) does not restrict the order of serial correlation in $\varepsilon_{i,j,t}$ (see Arellano (2003), p. 164). It is however reasonable to maintain that future realisations of idiosyncratic shocks do not affect the current relative wage; by contrast, we leave unrestricted the impact of past and current realisations of $\varepsilon_{i,j}$, as, for example, in the following general dynamic specification:

$$
\ln (w)_{i,j,t} = \sum_{s=1}^{k} \rho_s \ln (w)_{i,j,t-s} + \pi Z_{i,j,t} + \omega \varepsilon_{i,j,t} + u_{i,j,t} \quad (4.3)
$$

where $u_{i,j,t}$ is a white noise disturbance and $k$ is any suitable integer no smaller than unity.

Now, equations (4.2) and (4.3) together, besides implying $p \geq 1$, do have the implication that there should be no higher than $(p - 1)^{th}$ order serial correlation in $\varepsilon_{i,j}$, which can be tested empirically. The serial correlation tests reported in Table 2 bring no evidence for higher than first order serial correlation. This leads us to the conclusion that we could choose any $p$ no smaller than 2.

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25This approach is close to the specification analysis that is typically done in dynamic panel data models, which in addition to looking at overidentification statistics, searches indirect evidence on instruments validity by investigating the order of serial correlation in the error process. The only difference is that in dynamic panel data models the DGP of one of the predetermined explanatory variables, the lagged dependent, is known, whereas here we have to maintain a DGP like (4.3) for at least one predetermined or endogenous regressor, in order to obtain a testable restriction on the degree of serial correlation, given (4.2). In dynamic panel data models, however, DGP’s like (4.3) are implicitly maintained for the predetermined or endogenous explanatory variables of an unknown nature. For example, Bond (2002) suggests that "if $x_{it}$ is assumed to be endogenous then it is treated symmetrically with the dependent variable $y_{it}$. In this case the lagged values $x_{it-2}, x_{it-3} and longer lags (when observed) will be valid instrumental variables in the first-differenced equations for periods $t = 3, 4, ..., T$.” Bond’s suggestion boils down to maintain condition (4.2) with $p = 1$, or equivalently (4.3) and the lack of serial correlation. Notably, it has been incorporated by the endog option of the official Stata command for dynamic
We have considered several batteries of FODTSLS for various values of $p \geq 2$. Below, we report results for three different FODTSLS estimators, each using a different set of instruments and exogenous regressors, but all based on a "safe" choice of $p = 5$.

FODTSLS1 is a just identified estimator that maintains strict exogeneity of all $Z$ in addition to (4.2) and uses only one instrument given by the sixth lag of $\ln (w)_{i,j,t}$. Results for this estimator are presented in column 2 of Table 4. We note a remarkable increase in the first-stage F test, which was to be expected since $\ln (w)_{i,j,t-6}$ is more likely to capture the portion of cross-sectional variability in $\ln (w)_{i,j,t}$ that is due to different initial labour market conditions. Results change only slightly from the previous specification: in particular, relative wage and FDI in the Czech Republic lose significance. By contrast, results for FDI in the other two countries remain robust: the coefficient is again positive and significant for Poland, with point estimates remarkably close to those of column 1; for Hungary, instead, the estimated parameter on FDI is insignificant also in this case. Similarly, the exports parameter remains significantly negative and maintains nearly the same absolute size as before.

Up to now, results are based on the use of internal instruments only. Finding external instruments is generally quite complicated. Nevertheless, in a context like ours, one could try to use some measure of alternative wages for non-manual and manual employees at this purpose. For example, wages paid in non-manufacturing industries are likely to be at the same time uncorrelated with the disturbances, but correlated with the manufacturing wages; if this were so, one could use such wages as instruments for the relative wage. Based on these considerations, we use our previous measure of alternative relative wage, $\ln (w^o)_{j,t}$, as an instrument for the models, xtabond. The user-written Stata command xtabond2, written by Roodman (2006), while maintaining assumption (4.2), is more flexible allowing the user to choose a $p$ greater than 1.
log-relative wage, \( \ln (w)_{i,j,t} \). We maintain for \( \ln (w^o)_{j,t} \) the same sequential moment condition as for \( \ln (w)_{i,j,t} \). Notice that this approach is both consistent with, and more general than, the one used by Feenstra and Hanson (1997), who have to assume strict exogeneity of the alternative wages, in order to include them among the regressors. In column 3 of Table 4 we present results for FODTSLS2, a TSLS estimator with variables in forward orthogonal deviations and using an instrument set that consists of the sixth lag of the relative alternative wage, together with the same lag of the relative manufacturing wage and the fifth lag of real output. Now, the model is overidentified with two overidentifying restrictions; the Hansen J-statistics does not reject the validity of instruments at any conventional level of significance. Moreover, the F-statistics from first stage regression reassures about the relevance of these instruments. As before, results for FDI in Poland and for exports are not affected by these changes in the choice of instruments. As compared with column 2, relative wage and FDI in the Czech Republic gain significance, with the same signs as before. Point estimates remain almost completely unchanged.

Finally, in column 4, FODTSLS3 relaxes the strict exogeneity of output, maintained by all estimators described until now, to replace it with a sequential moment condition identical to (4.2). Now, the instrument set comprises selected GMM-type fifth and sixth lags of relative wage, relative alternative wage, and real output as instruments. The Hansen J-statistics supports the four overidentifying restrictions. The F-statistics from first stage regressions are inconclusive in this case, due to the presence of more than one endogenous variable. At the same time, however, the solution proposed by Stock and Yogo (2002) does not apply in our context, because it is designed for a spherical covariance matrix of the idiosyncratic error. Hence, we report F-statistics, even though we are aware of their inconclusiveness. With this caveat in mind, results appear extremely robust, parameter estimates maintaining the same sign and significance and
roughly the same absolute size as in column 1 and 3.

In order to further verify the robustness of our results, we also tried alternative lag structures for our instruments set; moreover, we also allowed the effects of real output to differ across countries. These alternative specifications did not brought about significant changes in the estimated parameters.

The presence of first-order autocorrelation in the residuals may be due to some form of dynamic misspecification. It is however well know (see, among others, Kiviet, 1995 and Bruno, 2005b), that standard $N$-consistent dynamic panel data estimators have poor finite sample performances with a small $N$. For this reason, we have applied the bias-corrected dynamic Within estimator as in Bruno (2005a,b); we have found no relevant changes both in sign and absolute size of the point estimates$^{26}$.

Above results obtained by means of TSLS strongly confirm our findings from the Within estimation. There is strong and robust evidence of heterogeneity in the impact of FDI on relative skilled labour demand. The contribution of international trade is always consistent with predictions from neoclassical trade theory.

In the second part of Table 4 we show results obtained by distinguishing the effect of exports across countries (MODEL 2). Columns follow the same order as those in the first half of the table, as far as the choice of instruments and estimation techniques are concerned. Overall, also in this case, TSLS confirm Within results. First, there is strong evidence supporting the imposition of the cross-country equality on the exports coefficients; the latter keep the same sign and significance as their Within counterparts (column 4 in Table 3), and point estimates are almost unaffected by the change in estimation method. Turning to $FDI$, cross-country heterogeneity is

$^{26}$All these estimates are available upon request.
strongly confirmed: in Poland, the effect is always positive and significant and point estimates are extremely close both to their fixed-effects analog and to their counterparts obtained under the assumption of constant exports coefficients. At the same time, the insignificant effect previously found for Hungary is strongly confirmed, as is the weak and non robust negative impact in the Czech Republic.

4.3. Quantifying the effects of FDI and international trade

Our results have confirmed theoretical predictions about the effects of FDI on relative skilled labour demand and wage inequality: the impact of FDI is likely to differ across countries, depending on their relative abundance of unskilled labour. Consistently with what emerges from other empirical studies (Feenstra and Hanson, 1997; Blonigen and Slaughter, 2001), we showed that FDI is more likely to exert positive effects, the more different the recipient country is from the investor in terms of relative abundance of unskilled labour. At the same time, we found robust evidence of a Stolper-Samuelson-type effect of trade in final goods: because these countries are relatively more unskilled abundant than their main trading partners, they tend to specialise in unskill-intensive productions; therefore, trade in final goods tends to shift relative skilled labour demand inward and to lower wage inequality.

But, how strong are these effects? In order to answer this question, we will now use estimated parameters from previous sections to compute skilled labour share elasticities to FDI and X. The formulas for these elasticities are:

$$\epsilon_{s,FDI} = \frac{\partial WSH_{i,j,t}}{\partial \ln FDI_{i,j,t}} \cdot \frac{1}{WSH_j} = \frac{\beta_{1,j}}{WSH_j}$$
and

\[ \epsilon_{s,X} = \frac{\partial WSH_{i,j,t}}{\partial \ln X_{i,j,t}} \cdot \frac{1}{WSH_j} = \frac{\beta_{2,j}}{WSH_j} \]

where \( WSH_j \) is the mean of \( WSH_j \) over the entire time period and across all industries in country \( j \). The name share elasticities highlights the fact that the above formulas represents only one part of the full expression for the elasticities of skilled labour demand: following Morrison and Siegel (2001), the latter would in fact be equal to:

\[
\eta_{s,FDI} = \epsilon_{s,FDI} + \xi_{C,FDI}
\]

and

\[
\eta_{s,X} = \epsilon_{s,X} + \xi_{C,X}
\]

where \( \xi_{C,FDI} \) and \( \xi_{C,X} \) are the cost elasticities to \( FDI \) and \( X \) respectively. These formulas suggest that the skilled labour demand elasticity to a given external factor is composed of two parts: the first accounts for the change in the marginal rate of technical substitution between skilled and unskilled labour induced by a variation in the external factor; the second, instead, accounts for the fact that a different value of the external factor corresponds to a different level of total costs, which in turn affects demand for skilled labour. However, since \( \xi_{C,FDI} \) and \( \xi_{C,X} \) cannot be identified without further information on the cost function, existing literature has relied only on share elasticities (Feenstra and Hanson, 1996a, 1997, 1999; Hijzen et al., 2005; 27Estimated share elasticities are industry and time specific. In order to save space, therefore, we will discuss share elasticities computed at the mean value of the skilled labor shares of total wage bill.

27
Crinò, 2006b), which provide a first piece of information about the effects of external factors like FDI and $X$ on relative labour demand for the skilled. In what follows, therefore, we will conform to this literature and use share elasticities with this interpretation.

Estimated share elasticities from MODEL 1 and MODEL 2 are reported in Table 5 together with their asymptotic standard errors; we also report share elasticities from MODEL 0, in order to ease comparison with other existing studies using a similar specification (Feenstra and Hanson, 1997; Lorentowicz et al., 2005). As already emerged, the effect of FDI on the skilled labour share of wage bill is positive and statistically significant in Poland: reported figures from MODEL 1 and MODEL 2 (our preferred specifications) imply that a 10% increase in FDI would raise the skilled labour share of wage bill by 0.36%-0.55%. In the Czech Republic, the same increase in FDI would instead result in a much smaller decline in the share: when significant, estimated share elasticities suggest that such a decline would range between 0.11% and 0.13%. In Hungary, instead, a change in FDI would not produce any significant variation in the skilled labour share of wage bill. Notice that the estimated share elasticities from MODEL 0 would be substantially higher, due to the use of a reduced-form specification of equation (4.1).

Turning to exports, estimated share elasticities from MODEL 1 and MODEL 2 suggest that a 10% increase in this variable would lower the skilled labour share of wage bill by about 1% in all the three countries. Also in this case, MODEL 0 would yield somewhat larger share elasticities.

Multiplying the estimated share elasticities by the change in FDI and exports between 1994 and 2002, as reported in Table 1, we can finally obtain an indication of the overall effect that each of these variables, alone, would have produced on the skilled labour share of wage bill in each country. These contributions are reported in the bottom panel of Table 5. As expected, the largest effect of FDI emerges for Poland: in this case, foreign penetration, alone, would have
increased the skilled labour share of wage bill by 10.5%-16%. These figures are somewhat lower that those reported by Lorentowicz et al. (2005) for the same country (34%) and by Feenstra and Hanson (1997) for Mexico (50%). This difference is likely to depend on the use of reduced-form equations - rather than structural equations - in those studies: if computed on share elasticities from MODEL 0, indeed, the contribution of FDI rises by about 10 percentage points and gets closer to that estimated by Lorentowicz et al. (2005). A second explanation for the difference is the use of a different proxy for foreign penetration in those studies: unlike us, in fact, both Lorentowicz et al. (2005) and Feenstra and Hanson (1997) use the fraction of foreign-owned firms on the total number of active firms. In any case, our results suggest that, even in the most conservative scenario, foreign penetration would have changed the earning distribution significantly in favor of Polish skilled workers. By contrast, estimated contributions for the Czech Republic confirm that, if anything, increasing FDI would have lowered only slightly the skilled labour share of wage bill, by a factor of about 3%. Turning to exports, contributions are similar both across countries and across specifications: increasing trade flows with more skilled abundant economies, alone, would have lowered the skilled labour share of wage bill by about 6% in Hungary, 4.8%-6.3% in the Czech Republic and 6%-7% in Poland.\footnote{We only compute contributions on country-specific exports semi-elasticities, since this makes it easier to understand the effect of exports, than computing contributions on the cross-country restricted share elasticities.}

5. Conclusions

In this paper, we studied the effects of inward FDI and trade in final goods on relative skilled labour demand and wage inequality in Poland, Hungary and the Czech Republic. The interest in these countries stems from the fact that, after the fall of the Communist regime in 1989, they
have undertaken a strong process of international integration - mainly with the neighboring EU Members - and have at the same time experienced a strong change in earning distribution in favor of the skilled.

We used an industry-year panel covering six manufacturing sectors between 1994 and 2002 and carried out a parametric analysis based on the estimation of a skilled wage-share equation derived from a short-run translog cost function. Estimation has been carried out through standard Within estimator, but the robustness of results has been assessed through several sensitivity tests based on different Instrumental Variable techniques.

Existing theoretical predictions and empirical results show that the effect of FDI on relative skilled labour demand is likely to vary across countries, especially depending on the relative skill abundance of the recipient economy as compared to the investor. Consistently, our estimates show strong heterogeneity in the FDI effect in the three economies: the effect is positive, significant and sizeable in Poland; on the contrary, it is always insignificant in Hungary; finally, FDI exerts at most a very small negative effect in the Czech Republic, even though such an effect is not always robust across specifications. These results reflect the relative position of the three countries as compared with the main investors, the EU Members: in particular, the positive effect in Poland seems to depend on the fact that this country has the largest relative endowment of unskilled labour among the three economies; these findings, indeed, are fully confirmed by other studies on the same country (Skuratowicz, 2001; Lorentowicz et al., 2005). As to trade in final goods, we find much more homogeneity in the results, which remain strikingly robust and significant across all specifications: in each and every country, increasing exports lower relative skilled labour demand, as predicted by standard neoclassical trade theories.

Our estimated share elasticities imply that foreign penetration, alone, would have increased
the skilled labour share of wage bill by 10.5%-16% in Poland, and decreased it by about 3% in the Czech Republic. Turning to exports, increasing trade flows, alone, would have lowered the skilled labour share of wage bill by about 6% in Hungary, 4.8%-6.3% in the Czech Republic and 6%-7% in Poland.

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**DATA APPENDIX**

Our dataset covers the period 1994-2002 and contains data on the following six ISIC Rev. 3 manufacturing industries: Food and tobacco; Textile, wood and paper; Petroleum, chemical,
rubber and plastic; Metal and mechanical products; Machinery, computers, RTV and communications; Vehicles and other transport equipment.

We gathered data on FDI stocks in Poland and Hungary from the OECD “International Direct Investment Statistics” (http://www.sourceoecd.org); for the Czech Republic, instead, we used the data on FDI inflows provided by the “Balance of Payments Statistics” of the Czech National Bank (www.cnb.cz) and retrieved FDI stocks as cumulative inflows.

Data on exports, imports and production come from the OECD “Stan Database for Industrial Analysis”, whereas data on total business expenditure on R&D are from the OECD “Basic Science and Technology Statistics” (http://www.sourceoecd.org).

Finally, we collected data on wages and employment of manual and non-manual workers from each country’s Statistical Yearbook. However, the Statistical Yearbooks of the Czech Republic do not report information on employment and wages of non-manual workers; therefore, we estimated them following Egger and Stehrer (2003): we first computed the number of non-manual employees as difference between total and manual employment, and then obtained the non-manual wage as $w^s_i = (N_i^s \bar{w}_i - N^u_i w^u_i)/N^s_i$, where the superscripts index skilled (non-manual) and unskilled (manual) workers, the subscript refers to the manufacturing industries, $N$ is the number of employees and $\bar{w}$ the average wage.

Moreover, the industrial breakdown of the employment and wage data is different from that of the remaining variables, since the former are classified in fourteen 2-digit NACE Rev.2 industries, whereas the latter follow the 2-digit ISIC Rev.3 classification. Hence, we converted the employment and wage data into the ISIC Rev.3 classification by means of the correspondence table provided by EUROSTAT (http://www.europa.eu.int/comm/eurostat). Finally, with the data in the new classification, we aggregated industries into the six ISIC sectors for which
we have data on FDI; the wage data for each of the six industries have been constructed as employment-weighed averages of the wages paid in the corresponding 2-digit ISIC Rev. 3 sectors. Table A1 reports the final correspondence between the FDI breakdown and the NACE Rev. 2 classification.

Real data (at constant 2000 prices) have been obtained by means of the GDP deflator provided by the OECD "Economic Outlook Database", which we used as common deflator across industries. PPP exchange rates from the OECD "Structural Statistics for Industry and Services" have been used to express all data in US$ for the parametric analysis.
COMPUTATIONAL APPENDIX

All computing work has been done in Stata SE version 9.2. The commands used for estimation are the following: *xtreg*, for Within; *ivreg2* and *xtivreg2* (Baum et al., 2003) for WTSLS; *ivreg2*, *xtivreg2* and *forthdev* (computing FOD’s and written by Bruno (2006)) for FODT-SLS. The bias approximation for WTSLS has been computed through Stata do-files written by the authors. For the bias-corrected dynamic Within estimator we have used *xtlsdvc*(Bruno, 2005b).

The clustered variance estimator has been computed through the option *cluster()* of *xtreg*, *ivreg2* and *xtivreg2*.

The serial correlation analysis has been based on *xtserial* (Drukker, 2003) for the Wooldridge test and *abar* (Roodman, 2006) for the AB tests.
References


