Margins of Multinational Labor Substitution

Employment at a multinational enterprise (MNE) responds to wages at the extensive margin, when an MNE enters a foreign location, and at the intensive margin, when an MNE operates existing affiliates. We present an MNE model and conditions for parametric and nonparametric identification. Prior studies rarely found wages to affect MNE employment. Our integrated approach documents salient labor substitution for German manufacturing MNEs and removes bias. In Central and Eastern Europe, most employment responds at the extensive margin, while in Western Europe the extensive margin accounts for around two-thirds of employment shifts. At distant locations, MNEs respond to wages only at the extensive margin.

Multinational enterprises (MNEs) are important mediators of world trade.[1] Surprisingly, however, their operation has rarely been found to strongly affect factor demands across locations.[2] We study how MNEs organize their global activity and the employment consequences at two critical margins. An MNE's labor demand responds to international wage differentials at the *extensive margin*, when the MNE enters a foreign market, and at the *intensive margin*, when the MNE operates existing affiliates.

Our empirical analysis shows that MNEs change their foreign presence only infrequently, but that these scant changes are associated with salient employment shifts. We devise an integrated estimation method to generalize earlier approaches in two important aspects. First, we estimate a system of labor demand outcomes in multiple locations at the intensive margin, including all relevant locations. Second, we jointly estimate presence in multiple foreign locations at the extensive margin. Rich data on German manufacturing MNEs and their majority-owned foreign manufacturing affiliates offer the required cost-function variables. Compared to single-equation estimation in the earlier literature, simultaneous labor-demand estimation for all relevant locations leads to substantively different results even in the absence of extensive-margin correction. In addition, results document that quantification of the extensive margin is economically and statistically important. In economic terms, we provide rigorous estimates of employment responses at the extensive margin. In Central and Eastern Europe, for example, where German MNEs expand most between 1996

[1]The world's ten largest MNEs in 2000 produce almost one percent of world GDP, and the one hundred largest MNEs are responsible for more than four percent of world GDP (UNCTAD press release TAD/INF/PR/47, 12/08/02). Transactions data for the United States show that around ninety percent of U.S. trade are conducted by firms with ownership of at least one related party abroad (Bernard, Jensen and Schott 2009).

[2]See e.g. Slaughter (2000) for U.S. and Konings and Murphy (2006) for European MNEs.

and 2001, our estimates suggest that the bulk of employment adjustment happens at the extensive margin. In Western Europe, in comparison, the extensive margin accounts for around two-thirds of overall employment shifts in response to German wages. In showing how MNEs first form and then operate their affiliate networks, our estimates provide labor-demand predictions to inform policy. In statistical terms, we document that omitting the extensive margin can bias intensive-margin estimates. An instance of complementarity bias can arise, for instance, if firms with a high likelihood to set up shop in low-wage locations also command comparatively low home wages, perhaps because their relocation propensity serves as a credible threat in wage bargaining at home.

At both margins, we find cross-wage elasticities with the home location to be strictly positive when statistically significant. So, home and foreign employment are substitutes within MNEs not only at the intensive but also at the extensive margin. Bootstrapped standard errors on the cross-wage elasticities reject equality between the intensive and the total elasticity of substitution for every location, corroborating the importance of the extensive margin. Elasticity point estimates at both margins are robust across different samples and wage data, model and correlation specifications, and parametric and nonparametric estimation techniques. With an eye on ease of implementation, we present conditions under which common Heckman (1979) correction and straightforward nonparametric estimation (similar to Das, Newey and Vella 2003) can be applied location by location and combined with outcome estimation—in our case a seemingly unrelated equation system of the MNE's labor demands. Beyond our context, the estimation technique potentially applies to empirical work on extensive margin adjustments more generally, such as export-market entry, import-market access, or intra-firm trade. Our generic MNE model unifies two strands of the existing empirical literature—one on MNE operations across existing locations and one on MNEs' location choices—into an integrated approach.[3]

For the intensive margin, Slaughter (2000) reports that MNE operations in low-wage locations have no detectable impact on relative home employment in U.S. industries. Similar to the U.S. evidence, Braconier and Ekholm (2000), Konings and Murphy (2006) and Marin (2004) find little or no evidence that operations of European MNEs in low-wage locations have an impact on home employment at the firm level. The latter papers estimate labor demand with independent regressions location by location. Using specifications similar to Konings and Murphy (2006), we

[3]Our generic MNE model encompasses motives for both vertical and horizontal foreign direct investment (FDI).

show for our MNE data that adopting a labor-demand system for all global locations leads to considerably different estimates even without extensive-margin correction. Recent evidence on U.S. MNEs by Harrison and McMillan (2006) suggests that foreign employment substitutes for U.S. MNE employment in industries with no significant intra-firm trade, whereas foreign employment in low-income countries complements U.S. employment in industries with strong intra-firm trade, so that the net employment effect is small at the intensive margin.[4] When an MNE relocates a production stage that is complementary at the intensive margin, however, job loss can still result at the extensive margin. Consistent with this idea, we find for distant overseas locations that the cross-wage elasticities are not statistically significant at the intensive margin but significantly positive at the extensive margin.

At the extensive margin, several earlier firm-level studies do not find wages or per-capita incomes to be significant predictors of location selection (e.g. Devereux and Griffith (1998) for U.S., Buch, Kleinert, Lipponer and Toubal (2005) for German MNEs).[5] Other studies, using multinomial logit estimation, find wages to predict location choice (e.g. Disdier and Mayer (2004) for French MNEs, and Becker, Ekholm, Jäckle and Muendler (2005) for Swedish MNEs and similar German MNE data as in this paper). But the multinomial logit model requires mutually exclusive location choices. So the simultaneous presence of MNEs in several locations needs to rest on the assumption of independent decisions, which is not necessarily compatible with MNE-wide profit maximization. In contrast, we condition on an MNE's past presence and its interaction with wages and find wage variables to be statistically significant predictors of location choice in selection regressions, while allowing for correlated decisions. Moreover, our MNE model shows that the full extensive-margin wage effect on employment is the product of two factors—the wage impact on the propensity of presence (as in the location-choice literature) times the impact of the propensity change on employment. Estimating the two factors together shows that wage differentials across locations are substantial predictors of an MNE's regional employments at the extensive margin.

[4] Riker and Brainard (1997) report too that affiliate activities in low-income countries are complementary to activities in high-income countries. Hanson, Mataloni and Slaughter (2005) shift focus from factor demands to intermediate input uses and report that affiliates of U.S. MNEs process significantly more intra-firm imports the lower are low-skilled wages.In contrast to the MNE evidence, Feenstra and Hanson (1999) attribute about a third of U.S. relative wage changes to offshoring at the sector level, within MNEs or across firms.

[5] Carr, Markusen and Maskus (2001) find evidence in aggregate data that relatively abundant high-skilled labor is a significant predictor of FDI of U.S. MNEs; and Blonigen, Davies and Head (2003) find that larger skill differentials predict less foreign MNE activity.

Our approach bridges the two literatures on employment substitution and location selection. There is a third MNE literature, which compares MNE performance to that of other MNEs or to national firms. Several studies detect no clear difference between MNEs and non-MNEs at the firm level (Egger and Pfaffermayr 2003, Barba Navaretti and Castellani 2008, Jäckle and Wamser forthcoming), with few exceptions (Desai, Foley and Hines 2009, Debaere, Lee and Lee 2006). Allowing output to vary, Desai et al. (2009) find that foreign and domestic investment expenditures and wage bills are positively associated within MNEs. At the worker level, Becker and Muendler (2008) document more worker retentions at expanding MNEs than at non-expanding MNEs. Those comparisons leave unanswered, however, whether an MNE's gain in market shares after cost-saving FDI comes at the expense of national competitors' market shares. The method in this paper intentionally holds MNEs' market shares constant, by conditioning on output as cost function estimation requires. So identification of employment shifts within MNEs is consistent with multiple forms of product-market competition under lean assumptions. The analysis of employment shifts between MNEs and their competitors after FDI, in contrast, would require additional data on domestic competitors, and additional assumptions on product market structure, the degree of competition, and the elasticity of product demand. It remains an open task for research to discern whether foreign MNE expansions stabilize industry employment at home or whether market-share gains by MNEs result in market-share losses at national competitors.[6]

This paper has four more sections. In Section I, we present a model of the expansion and operation of MNEs and report identification conditions for estimation under location selectivity (derivations in the Appendix). Section II discusses the data and descriptive statistics on location choice (details in the Appendix). Estimates of multinational labor substitution are presented in Section III, and interpreted in counterfactual evaluations. Section IV concludes.

[6] Preliminary reduced-form regression evidence from German plant-level data linked to MNE data from this paper suggests that foreign employment at MNEs is also associated with reduced competitor employment, consistent with a market-stealing effect of expanding MNEs from their competitors.

I Multinational Expansion and Operation

A Labor demand and location selectivity

There are L locations for production and sales. In period t, MNE j employs $\mathbf{y}_{jt}$ workers at up to L locations and produces up to L location-specific outputs $\mathbf{q}_{jt}$ with quasi-fixed capital $\mathbf{k}_{jt}$ under variable-input prices $\mathbf{w}_t$ (these variables are L-dimensional vectors). Production technology is the same for all MNEs. The factor prices $\mathbf{w}_t$ are market-wide outside prices by location. We specify the short-run cost function $C(\mathbf{q}_{jt}; \mathbf{k}_{jt}, \mathbf{w}_t)$ to be a multiproduct translog cost function. The translog form is flexible. Its cross-wage elasticities of substitution offer a compact way to summarize multinational labor substitutability or complementarity.[7]

An MNE's wage bill shares are $s^\ell_{jt} \equiv w^\ell_t y^\ell_{jt}/C_{jt}$ at locations $\ell = 1, \ldots, L$. Under a translog short-run cost function, we can transform the L equations of wage-bill shares into L labor demand functions by multiplying the dependent variable and all regressors with the observation-specific scalars C_{jt}/w^ℓ_t and obtain the following labor demands $y^\ell_{jt} = \partial C_{jt}/\partial w^\ell_t = s^\ell_{jt} C_{jt}/w^\ell_t$:

$$y^\ell_{jt} = \mathbf{x}^\ell_{jt}\beta^\ell + \epsilon^\ell_{jt} \qquad\qquad (\ell = 2, \ldots, L), \tag{1}$$

where

$$\mathbf{x}^\ell_{jt}\beta^\ell = \alpha_\ell \tfrac{C_{jt}}{w^\ell_t} + \sum_{n=1}^{L} \left(\mu_{\ell n} \ln\!\left[(q^n_{jt})^{C_{jt}/w^\ell_t}\right] + \kappa_{\ell n} \ln\!\left[(k^n_{jt})^{C_{jt}/w^\ell_t}\right] + \delta_{\ell n} \ln\!\left[(w^n_t)^{C_{jt}/w^\ell_t}\right] \right)$$

by Shepard's lemma (see Appendix A).[8]

Not all firms produce at all locations. The employment effect of MNE selection into locations is both of economic interest and of statistical relevance for estimating (1). In economic terms, our approach permits estimation of cross-wage elasticities of labor demand at the extensive margin. In terms of statistical properties, potential selection bias for intensive-margin elasticities is

[7] We choose a short-run function because our location-selectivity estimation captures long-term installation costs and because observed capital inputs are arguably closer proxies to MNE-specific user costs of capital than price measures. We use time subscripts to clarify that our empirical approach compares firm j's current presence to its own past presence, requiring panel data.

[8] The transformed labor-demand equations have three advantages over conventional wage-bill share equations. First, labor demand is not bounded above so that, conditional on $\mathbf{x}^\ell_{jt}$, the labor demand disturbance satisfies the assumption of one-sided censoring for selectivity correction. Second, wages become regressors only and do not enter the dependent variable. Third, there is no constant term among the regressors $\mathbf{x}^\ell_{jt}$ so that lacking identification of the constant in a nonparametric selection correction is no concern.

treated with extensive-margin controls.[9] Previous empirical research largely ignored selection in the context of multinational labor demand.

An MNE's choice of foreign activity can be understood as a two-stage decision. At time $t-\tau$, that is τ periods prior to production and sales, the MNE selects the locations for its foreign affiliates and capital inputs $\mathbf{k}_{jt}$ around the world. The MNE faces uncertainty and bases the location and capital-input decisions on the vector of selection predictors $\mathbf{z}_{j,t-\tau}$ (competitors' future outputs $\mathbf{q}_{i\neq j,t}$, own realized output $\mathbf{q}_{jt}$ and input prices $\mathbf{w}_t$ are uncertain). On the second stage at time t, MNE j simultaneously chooses output $\mathbf{q}_{jt}$ and variable factor inputs. So, conditional on presence $d^\ell_{j,t}=1$ at location ℓ, the expectation of observed MNE employment $\bar{y}^\ell_{jt}$ is

$$\bar{y}^\ell_{jt} = \mathbb{E}\left[\mathbf{x}^\ell_{jt}\beta^\ell + \epsilon^\ell_{jt}\,|\,\mathbf{x}^\ell_{jt}, \mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}\right] = \mathbf{x}^\ell_{jt}\beta^\ell + m^\ell(\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau})) \tag{2}$$

by (1), where $\mathbf{d}_{jt}$ is a vector of MNE j's multinational presence at locations $n = 1, \ldots, L$ and $\mathbf{P}_{jt}$ is a vector of propensities $\Pr^\ell_{jt}(\mathbf{z}_{j,t-\tau}) = \mathbb{E}\left[d^\ell_{jt}|\,\mathbf{z}_{j,t-\tau}\right]$ for MNE j to be present at locations $n = 1, \ldots, L$. The empirical concern is that eq. (2) violates mean independence of the disturbance if the selectivity term $m^\ell(\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau})) = \mathbb{E}\left[\epsilon^\ell_{jt}|\, d^\ell_{jt}=1;\, \mathbf{z}_{j,t-\tau}\right] \neq 0$.

In economic terms, permanent wage differentials between locations ℓ and n have an impact on labor demand at two distinct margins. At the *intensive margin*, the spot wage w^n_t affects expected employment $\bar{y}^\ell_{jt}$ through $\mathbf{x}^\ell_{jt}\beta^\ell$ (regressors $\mathbf{x}^\ell_{jt}$ include the translog-transformed spot wages w^n_t). We call this the intensive-margin response because the spot wage affects employment outcomes conditional on the MNE's presence throughout the world. At the *extensive margin*, past wages affect a firm's propensity $\Pr^n_{jt}$ to enter n, and in turn presence at n affects current employment at ℓ through $m^\ell(\cdot)$ (selection predictors $\mathbf{z}_{j,t-\tau}$ contain past wages $w^n_{t-\tau}$). Note that, in our context

[9] Consider, for instance, the effect of home wages (n = HOM) on employment in Central and Eastern Europe (ℓ = CEE). In the absence of selectivity treatment, the CEE wage-bill response to log home wages is measured by $\delta_{\ell n}$, and a positive $\delta_{\ell n}$ implies substitutability between home and CEE employment; a negative $\delta_{\ell n}$ is necessary for complementarity. Suppose German firms that face high wages under an industry-specific collective agreement also have a high likelihood to set up shop in CEE countries. For such firms, the uncorrected estimate of the CEE wage-bill response to home wages is positively biased so that the estimated cross-wage elasticity will be biased towards substitutability between home and CEE countries, unless selectivity is controlled for. Such a substitutability bias is plausible if cost uncertainty in the host location, or an industry's low relocation propensity in the past, make relocation a weak threat with little credibility in wage bargaining at home. For other foreign locations, the threat may be more credible. Suppose that firms with a high likelihood to set up shop in developing countries (DEV) command comparatively low home wages because their relocation propensity, given large labor-cost differences, serves as a credible threat in wage bargaining at home. For such firms, the uncorrected estimate of the DEV wage-bill response to home wages is negatively biased so that the estimated cross-wage elasticity is biased towards complementarity between home and DEV countries, unless selectivity is controlled for.

of cross-location employment responses, the extensive margin cannot be represented with just a count of affiliates or employment because the opening of affiliates has an unobserved effect on MNE employment elsewhere. Our structural estimation accounts for such employment shifts across locations at the extensive margin.

A permanent wage change at location n results in an overall labor-demand response at location ℓ by

$$\frac{\partial \bar{y}^{\ell}_{jt}}{\partial w^{n}} = \frac{\partial y^{\text{int},\ell}_{jt}}{\partial w^{n}_{t}} + \frac{\partial y^{\text{ext},\ell}_{jt}}{\partial \Pr^{\ell}_{jt}} \cdot \frac{\partial \Pr^{\ell}_{jt}}{\partial w^{n}_{t-\tau}}, \tag{3}$$

where $y^{\text{int},\ell}_{jt} \equiv \mathbf{x}^{\ell}_{jt}\beta^{\ell}$ and $y^{\text{ext},\ell}_{jt} \equiv m^{\ell}(\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau}))$.[10]

Estimators for one margin at a time can fail to detect the correct magnitude of employment responses to international wage differences for at least two reasons. First, at the intensive margin, eq. (2) shows that β^{ℓ} coefficients may be biased unless the unobserved error component $m^{\ell}(\cdot)$ is controlled for. Bias arises if the extensive-margin term is omitted but correlated with the intensive-margin term. A preview of results documents bias (Table 9). Several uncorrected cross-wage elasticities are distorted towards complementarity (negative relative differences in the table) and can even turn from indicating substitutability to complementarity in several cases (relative differences of less than negative one in the table). But the distortion is not uniform across foreign locations. The home-employment elasticity with respect to foreign wages, for instance, is under-estimated with a complementarity bias of up to 270 percent for wage changes in developing countries (DEV) and over-estimated with a substitutability bias of up to 15 percent for Central and Eastern European (CEE) wage changes at the intensive margin.[11]

Second, eq. (3) clarifies that the extensive margin response is the product of two factors: the effect of wages on the propensity of presence at a location ($\partial \Pr^{\ell}_{jt}/\partial w^{n}_{t-\tau}$), multiplied by the effect of the presence propensity on employment ($\partial y^{\text{ext},\ell}_{jt}/\partial \Pr^{\ell}_{jt}$). The wage effect on presence propensity has been estimated in the earlier literature on location choice but, by eq. (3), it measures only

[10] In general, the overall labor-demand response at location ℓ is

$$\frac{\partial \bar{y}^{\ell}_{jt}}{\partial w^{n}} = \frac{\partial y^{\text{int},\ell}_{jt}}{\partial w^{n}_{t}} + \sum_{k=1}^{L} \frac{\partial y^{\text{ext},\ell}_{jt}}{\partial \Pr^{k}_{jt}} \cdot \frac{\partial \Pr^{k}_{jt}}{\partial w^{n}_{t-\tau}}.$$

We treat the general case with nonparametric estimation. Under parametric location choice (Heckman 1979), the overall labor-demand response simplifies to (3).

[11] Both the uncorrected and the selectivity-corrected estimate are statistically significant at the 5-percent level for CEE, but not statistically significant for DEV. There are six cases of complementarity bias for statistically significant uncorrected and selectivity-corrected cross-wage elasticities.

one part of the extensive margin's importance for employment outcomes. Similar to earlier findings, our data exhibit only a weak association between wages and the presence propensity, that is $\partial \Pr_{jt}^{\ell}/\partial w_{t-\tau}^{n}$ is small. This is consistent with sunk costs that make extensive-margin adjustments infrequent and hard to measure. But, once appropriately weighted with the associated employment response $\partial y_{jt}^{\text{ext},\ell}/\partial \Pr_{jt}^{\ell}$, wage changes at the extensive margin are found to have an economically and statistically significant impact.

B Elasticities

Cost-function estimates themselves are hard to interpret. We therefore report results in terms of cross-wage elasticities of substitution. These elasticities quantify the response of labor demand in one location to permanent wage changes at the same location or elsewhere. Our model of the MNE allows us to express constant-output cross-wage elasticity of substitution between factors ℓ and n.[12] The *cross-wage elasticity of substitution* is defined as $\varepsilon_{\ell n} \equiv \partial \ln y_{jt}^{\ell}/\partial \ln w^{n}$ and becomes

$$\varepsilon_{\ell n}^{T} = \frac{\partial s_{jt}^{\ell}/\partial \ln w^{n}}{s^{\ell}} + s^{n} \quad (n \neq \ell) \qquad \text{and} \qquad \varepsilon_{\ell\ell}^{T} = \frac{\partial s_{jt}^{\ell}/\partial \ln w^{\ell}}{s^{\ell}} + s^{\ell} - 1 \tag{4}$$

for a short-run translog cost function, where $s^{\ell} = w^{\ell} y^{\ell}/C_{jt}$ is the wage bill share of location ℓ in the MNE's total wage bill. By (2) and 3, the marginal response of the wage bill share s_{jt}^{ℓ} to a permanent change in $\ln w^{n}$ is

$$\partial s_{jt}^{\ell}/\partial \ln w^{n} = \delta_{\ell n} + \frac{\partial \mathbb{E}\left[\epsilon_{jt}^{\ell} \mid \cdot\right]}{\partial w_{t-\tau}^{n}} \frac{w_{t}^{\ell} w_{t}^{n}}{C_{jt}} = \delta_{\ell n} + \frac{\partial m^{\ell}(\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau}))}{\partial \mathrm{Pr}_{jt}^{\ell}} \frac{\partial \mathrm{Pr}_{jt}^{\ell}}{\partial w_{t-\tau}^{n}} \frac{w_{t}^{\ell} w_{t}^{n}}{C_{jt}}. \tag{5}$$

The cross-wage elasticities draw on three sources of information. The first term in (5) measures the labor demand response at the intensive margin, where $\delta_{\ell n}$ is the regression coefficient on the transformed wage in labor demand. The second term in (5) measures the employment response at the extensive margin. As we demonstrate below, the second term can be inferred from location-choice coefficients and labor-demand coefficients.[13] Third, observed wage-bill shares enter in (4). If there is no response of the wage bill share to a marginal log wage change ($\partial s_{jt}^{\ell}/\partial \ln w^{n} = 0$),

[12]The cross-wage elasticity provides the same information to determine complementarity and substitutability as the Allen-Uzawa elasticity, which scales the cross-wage elasticity by a cost share. The Morishima elasticity measures curvature but is less informative regarding complementarity.

[13]The extensive-margin estimate is multiplied by the spot wage w_{t}^{n} because location-choice estimation uses w_{t}^{n} as regressors, not their logs. Division by C_{jt}/w_{t}^{ℓ} converts the extensive-margin estimate from the transformed labor demand eq. (1) back into the wage bill share equivalents.

then the cross-wage elasticity in (4) simplifies to s^n, the wage bill share at the location where the wage hypothetically varies. The reason is that, if there is no response of the wage bill share to the wage, then employment must vary in an exactly offsetting way against the wage.

The cross-wage elasticities are constant-output elasticities and reflect the curvature of the firm's multinational production technology. For the estimation of (2), we therefore condition on the vector of location-specific outputs. In product-market equilibrium, of course, an MNE's market share is endogenous to its cost or sales advantages after FDI. As mentioned, this suggests an extended approach with endogenous output for future research.

C Modelling selectivity

A profit-maximizing firm is present at location ℓ iff the expected profit difference between presence and absence strictly exceeds the sunk costs of presence:

$$\begin{aligned} d_{jt}^{\ell} &= \mathbf{1}\Big(\mathbb{E}_{j,t-\tau}[p^{\ell}q_{jt}^{\ell,*}] + \mathbb{E}_{j,t-\tau}[C(q_{jt}^{\ell}{=}0;\cdot) - C(q_{jt}^{\ell,*};\cdot)] - F_{j,t-\tau}^{\ell} + \eta_{j,t-\tau}^{\ell} > 0\Big) \\ &= \mathbf{1}\left(H(\mathbf{z}_{j,t-\tau}) + \eta_{j,t-\tau}^{\ell} > 0\right), \end{aligned} \tag{6}$$

where $F_{j,t-\tau}^{\ell}$ is the sunk cost of producing at ℓ, and $\eta_{j,t-\tau}^{\ell}$ is an MNE's specific disturbance to sunk costs. The expected net profit of presence $H(\mathbf{z}_{j,t-\tau})$ is equal to the sum of expected revenues $p^{\ell}q_{jt}^{\ell,*}$ from producing at ℓ and the expected cost savings $C(q_{jt}^{\ell}{=}0;\cdot) - C(q_{jt}^{\ell,*};\cdot)$ from presence at ℓ, less the sunk cost.

Our empirical MNE model has $L-1$ location-selection equations (6) because presence at home cannot be estimated in a data set for a single country's MNEs. The model has $L-1$ outcome equations (2) because the cost function is homogeneous of degree one in wages and hence one labor-demand equation becomes redundant (we omit the home labor-demand equation). Denoting home with $\ell{=}1$, the estimation model is therefore

$$\begin{aligned} d_{jt}^{\ell} &= \mathbf{1}\left(H(\mathbf{z}_{j,t-\tau}) + \eta_{j,t-\tau}^{\ell} > 0\right), \qquad (\ell = 2,\ldots,L) \\ y_{jt}^{\ell} &= \mathbf{x}_{jt}^{\ell}\beta^{\ell} + m^{\ell}(\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau})) + \epsilon_{jt}^{\ell} \end{aligned}$$

by (6) and (2). Different functional forms can be specified for $H(\mathbf{z}_{j,t-\tau})$, and alternative distributional assumptions can be placed on $\eta_{j,t-\tau}^{\ell}$ and ϵ_{jt}^{ℓ}. We consider two sets of assumptions: (1) a parametric version with linear $H(\cdot)$ and joint normality of $\eta_{j,t-\tau}^{\ell}$ and ϵ_{jt}^{ℓ}; (2) a nonparametric version for some smooth function $H(\cdot)$ with independent $\eta_{j,t-\tau}^{n}$ and ϵ_{jt}^{ℓ} for $n \neq \ell$.

Assumption A: Parametric location selection. Assumption (1) is an extension of the familiar Heckman (1979) selection model to multiple equations (locations). The correlation between ϵ^n_{jt}, the idiosyncratic component of labor demand, and $\eta^\ell_{j,t-\tau}$, the unobserved labor-demand effect of location selectivity, across locations $n \neq \ell$ is crucial for estimation of outcomes (2). Our data reject independence of ϵ^n_{jt} and $\eta^\ell_{j,t-\tau}$.[14] To specify a correlation structure consistent with these findings, we depart from the idea that selection disturbances include both location-specific parts such as, for example, surprising changes to profit repatriation policies in the host country and include MNE-specific parts such as idiosyncratic shocks to a firm's sunk entry costs. Changes to host-country repatriation policies affect the entry decision. Once the MNE operates in the host country, it minimizes costs irrespective of entry-related host-country shocks. So, we consider it plausible to assume that there is an MNE-specific, location-independent component e_{jt} to the selection shock $\eta^n_{j,t-\tau}$ and that the labor-demand shock ϵ^ℓ_{jt} correlates with the selection shock $\eta^n_{j,t-\tau}$ elsewhere only through the MNE-specific component e_{jt}. The assumption is not rejected in our data. Note that, under this assumption, cost function disturbances do covary with entry shocks across locations, but only through an MNE-specific component.

Lemma 1 in Appendix C shows that under this assumption, location-by-location correction for selectivity is permissible. Intuitively, all selection-related information that is relevant for labor demand at any location ℓ is fully contained in the single presence indicator d^ℓ_{jt}, which is as informative about $\eta^\ell_{j,t-\tau}$ as any other location indicator.

Assumption B: Nonparametric location selection. Under nonparametric selectivity correction, no functional-form assumption needs to be placed on the distributions of $\eta^\ell_{j,t-\tau}$ or ϵ_{jt}, and $H(\cdot)$ can be any smooth function. We consider a nonparametric multiple-outcome model with multiple thresholds. We present assumptions that guarantee identification similar to a single-outcome model with multiple thresholds in Das et al. (2003). A set of sufficient identifying assumptions is stated in Appendix D, where we also provide a proof (Lemma 2) that transforms a related result from Das et al. (2003) to our context.

We base identification on four sufficient conditions. First, the conditional expectation of the labor demand disturbance $\eta^\ell_{j,t-\tau}$ is a differentiable function of propensity scores. Second, at least

[14] SUR estimation of the outcome equations shows that ϵ^n_{jt} and ϵ^ℓ_{jt} correlate so that ϵ^n_{jt} and $\eta^\ell_{j,t-\tau}$ must be correlated because ϵ^ℓ_{jt} and $\eta^\ell_{j,t-\tau}$ are correlated.

one predictor of the propensity score is not also a predictor of the labor-demand outcome. Third, the regressors in the information set at $t - \tau$ predict the propensity score. Note that these three conditions allow us to relax the earlier identifying assumption that $(\epsilon^n_{jt}, \eta^\ell_{j,t-\tau})$ is independent of $\mathbf{x}^m_{jt}$ and $\mathbf{z}_{j,t-\tau}$ for all ℓ, m, n. Compared to Assumption A, these three assumptions only require that, conditional on the propensity score Pr^ℓ_{jt}, ϵ^ℓ_{jt} is uncorrelated with all functions of $\mathbf{x}^\ell_{jt}$ and $\mathbf{z}_{j,t-\tau}$. Fourth, we impose cross-equation independence on the labor demand disturbance $\eta^\ell_{j,t-\tau}$ (so that we do not need to condition on observed $\mathbf{d}^{k\neq\ell}_{jt}$ elsewhere). The nonparametric estimator allows for conditional heteroskedasticity of unknown form (and thus presents a nonparametric alternative to Chen and Khan's (2003) three-step estimator). This makes nonparametric analysis a powerful tool for multivariate binary selection estimation.

D Modelling wage endogeneity

We derived cross-wage elasticities in the benchmark context of competitive labor markets above. MNEs, however, are known to pay wage premia over local competitors. Suggested reasons include a more highly skilled workforce composition and rent sharing through efficiency wages or bargaining. We address this concern both theoretically and empirically in Appendix E. We show that cross-wage elasticities (5) are consistent with departures from competitive labor markets under wage bargaining such as in Stole and Zwiebel (1996a, 1996b). To mitigate empirical concerns with endogenous wage premia, we predict the log wage residual ψ^ℓ_{jt} from a reduced form log wage regression $\ln w^\ell_{jt} = W(\mathbf{z}_{j,t-\tau}) + \psi^\ell_{jt}$, where $W(\cdot)$ mirrors the functional form of the location-selection equation. We then include the predicted residual wages ψ^ℓ_{jt} for all locations with an MNE's presence as controls in outcome eq. (2) on the second stage.[15] This addresses concerns with profit-related pay (profitability may covary with industry-median wages) and unobserved workforce heterogeneity (unobserved MNE productivity could bias our employment estimation). The predicted wage residuals are orthogonal to the propensities of location selection by construction and serve as controls only. In line with the structural MNE model, estimation of cross-wage elasticities rests solely on the industry-wide median wage coefficients ($\delta_{\ell n}$).

[15] We thank an anonymous referee for this suggestion.

II Data and Descriptive Statistics

Our principal data source is a confidential three-dimensional panel data set of German MNEs (parent-affiliate-year observations), collected by Deutsche Bundesbank (BuBa). Individually identified outward FDI data are available since 1996, include all directly and indirectly owned foreign affiliates above reporting thresholds, and provide two-digit NACE 1.1 sector classifications for the parent and affiliates. Our estimation sample ends in 2001.

We retain only majority-owned affiliates because a multi-location cost function suggests that parent firms have full managerial control.[16] We restrict the sample to manufacturing parents and their manufacturing affiliates. MNEs that span fewer industries appear more likely to satisfy the assumption of full managerial control, and cross-country wage data are most comprehensive and reliable for the manufacturing sector. Results for majority-owned affiliates from any sector (and their manufacturing parents) are nevertheless broadly similar.[17]

We transform the data to parent-location-year observations, deflate them with location-specific CPIs, convert foreign-currency values to their EUR equivalents in December 1998 (the sample mid point) to remove nominal exchange rate fluctuations, and combine the data with complementary information on wages and host-country characteristics from various sources. Details on currency conversion and the complementary host-country data are in Appendix F.

A MNE data

For foreign affiliates, we obtain employment, turnover and fixed assets from BuBa's MIDI database (MIcro database Direct Investment, formerly DIREK). MIDI covers the universe of majority-owned foreign affiliates and offers their balance sheet information, including in years with zero turnover. MIDI is based on outward FDI information from a legally mandated annual survey that covers

[16]Majority ownership has the additional advantage to be insensitive to a change in the reporting threshold in MIDI 1999. German parent firms may in turn be ultimately owned by foreign MNEs; between 1996 and 2001 13.1 percent of the German MNEs in our sample are affiliates of foreign MNEs.

[17]Employment at non-manufacturing affiliates abroad is important. Majority-owned retail and wholesale affiliates of manufacturing parents, for instance, account for about as much employment abroad as majority-owned manufacturing affiliates worldwide (but in Central and Eastern Europe for just about half as much employment as manufacturing affiliates). In a sample with majority-owned affiliates from any sector (and their manufacturing parents), labor substitution at both margins is even more pronounced than in our manufacturing-affiliate sample, while the intensive margin becomes relatively more important perhaps because of lower sunk entry and exit costs outside manufacturing. Absent selectivity correction, distortions into complementarity are more frequent in the sample with affiliates from any sector.

the universe of German parent firms with foreign corporate holdings above minimum ownership shares and capital stock thresholds (Lipponer 2003). We use fixed assets from the balance sheet as our measure of the capital stock, thus excluding non-physical capital to avert valuation differences across firms. Turnover is not consolidated for within-MNE shipments, but is a proxy nevertheless to affiliate production for cost-function estimation.

For German parent firms, employment, turnover and fixed assets come from BuBa's confidential USTAN database (Deutsche Bundesbank 1998), which records balance sheets and income statements of firms that draw a bill of exchange. The bill of exchange is a common form of payment among firms of all sizes throughout the sample period 1996-2001 (though losing popularity thereafter). USTAN is considered the most comprehensive source of balance sheet data for companies of all sizes outside the financial sector in Germany. We link MIDI and USTAN data by parent name and address, resulting in the loss of some observations from the universe.[18] From USTAN, we retain non-MNEs (national firms) that are to become MNEs during the sample period or were MNEs earlier in the sample period.

To reduce dimensionality, we lump host countries into four *aggregate locations*: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), and WEU (Western Europe), beyond the home location Germany (see Table G.2 for definitions). Aggregation into four foreign locations and home limits the estimated cross-wage elasticity matrix to five columns and rows (with 25 elasticity estimates). We choose the aggregate locations to share geographic characteristics, and to broadly contain countries with relatively similarly skilled labor forces or related institutional characteristics. CEE and WEU share borders with Germany and are geographically contiguous, whereas OIN includes non-European industrialized countries, and DEV spans the remaining developing countries throughout Africa, Latin America and the Asia-Pacific region. The aggregate locations, especially DEV, may conceal considerable heterogeneity so that we emphasize patterns in results that apply to more than one location and conduct robustness estimation grouping countries into four manufacturing-wage quartiles.

As Table 1 shows, the four aggregate locations host similarly large manufacturing workforces for German manufacturing MNEs: between 250,000 and 400,000 employees. Among the low-

[18]Our conservative string matching routine filters out potential duplicates from time-varying firm identifiers in USTAN. In manual treatments, only doubtlessly identifiable parent pairs from MIDI and USTAN are kept. At the expense of reduced sample size, this caution guarantees the formation of time-consistent parent pairs.

Table 1: MNEs and Labor Markets

	HOM	CEE	DEV	OIN	WEU
	(1)	(2)	(3)	(4)	(5)
MNE employment	1,423,086[a]	245,721	332,622	319,221	394,579
Estimation sample MNE employment	962,726	125,199	184,560	139,240	191,854
Mean employment per sample MNE	1,629.0	387.6	407.4	736.7	282.6
Individual affiliates' employment share		.0003	.0001	.00008	.0002
All German MNEs' total employment share	.175[a]	.014	.002	.006	.021
MNE log wage premia over local competitors[b]		.645	.855	.094	.288

Sources: MIDI and USTAN 2000 (1996 to 2001 for prediction), German manufacturing MNEs and their majority-owned foreign manufacturing affiliates; ILO paid manufacturing employment by country in 2000; UNIDO manufacturing wages 1998 and IUI 1998 paid wages at majority-owned manufacturing affiliates of Swedish manufacturing MNEs.

Notes: Employment shares are location-wide averages over country-mean shares for affiliates of German MNEs in ILO totals. Wage premia are logs of the ratios of paid wages at Swedish MNEs over UNIDO manufacturing wages. Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

[a]Predicted German employment at in- and out-of-sample MNEs, based on linear employment regressions to account for incomplete MIDI-USTAN matches.

[b]Swedish MNE log wage premia. German and Swedish MNEs exhibit similar labor-demand behavior (Becker et al. 2005).

wage locations we focus on CEE where most expansions happen. For the 2,247 MIDI MNEs with foreign presence either in 1996 or 2000, CEE was the region where MNEs opened most new affiliates, operating 18.2 percent more affiliates in 2000 than in 1996, followed by DEV with a 12.6 percent increase, OIN with 3.2 percent and WEU with 2.0 percent. We estimate that German manufacturing MNEs with majority-owned foreign manufacturing affiliates employ about 1.4 million German workers in 2000, including their predicted out-of-sample employment.[19] The largest employment per MNE occurs in OIN and the smallest employment in WEU.

Table 1 also presents a comparison of German MNE employment figures to ILO employment totals. Although German manufacturing MNEs employ an estimated 17.5 percent of German manufacturing workers, the MNEs' domestic labor-market power is arguably limited. In Germany, collective agreements with industry-wide unions do not allow member firms in the employer association to deviate from wage schedules that specify wages by worker skill and seniority (but distressed member firms qualify for exception clauses and non-member firms are not bound). Ger-

[19]MIDI and USTAN matches are incomplete so that we do not observe parent employment for every German MNE. We predict total parent employment for the full sample of German manufacturing MNEs from a linear regression of parent employment on foreign employments.

Table 2: LOCATION COUNTS BY MNE

			L in 2000			*Total*
L in 1996	1	2	3	4	5	(100%)
1	*0.0%*	*83.5%*	*12.2%*	*2.6%*	*1.6%*	*794*
2		83.7%	12.5%	3.2%	0.6%	687
	34.7%	*54.7%*	*8.2%*	*2.1%*	*0.4%*	*1,052*
3		23.7%	55.8%	15.8%	4.7%	190
	28.0%	*17.1%*	*40.2%*	*11.4%*	*3.4%*	*264*
4		11.1%	25.0%	45.8%	18.1%	72
	24.2%	*8.4%*	*19.0%*	*34.7%*	*13.7%*	*95*
5		7.4%	3.7%	22.2%	66.7%	27
	35.7%	*4.8%*	*2.4%*	*14.3%*	*42.9%*	*42*
Total		630	211	91	44	976
	477	*1,293*	*308*	*112*	*57*	*2,247*

Source: MIDI universe 1996 and 2000 (not matched to USTAN), manufacturing MNEs and their majority-owned foreign manufacturing affiliates.
Notes: MNEs with foreign presence in 1996 and 2000 (large entries), and MNEs with foreign presence in one or both years (small entries). Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

man MNE affiliates have plausibly small market power abroad. Individual foreign affiliates of German MNEs command an average market share of just between .7 percent of a percent (.00007 in OIN) and 3 percent of a percent (.0003 in CEE) across the four foreign regions; even the total of all German MNEs merely commands an average market share by foreign country of between .2 percent (DEV) and 2.1 percent (WEU). Observed wage premia suggest that MNEs do not exert monopsony power. Canonical monopsony models predict wage mark-downs. In contrast, MNEs pay wage premia (Swedish MNEs pay between 7.2 percent (OIN) and 86.1 percent (DEV) over their local competitors). To control for potential labor-demand distortions from MNE rent sharing or unobserved skill compositions, we account for MNE wage premia at home and abroad in estimation.

The data exhibit strikingly rare changes to foreign presence, consistent with considerable sunk costs of entry and exit. Table 2 shows changes to foreign presence between 1996 and 2000. Large-font entries are for firms that are MNEs in both years, italicized small-font entries include firms that become or cease to be MNEs. Out of five MNEs with two locations (home and one foreign

Table 3: MNE COUNTS OF CHANGING AFFILIATE NUMBERS

	CEE	DEV	OIN	WEU	*MNE Total*
$N_{2000} - N_{1996}$	(1)	(2)	(3)	(4)	(5)
≤ -3	2	3	2	15	22
-2	3	11	3	14	31
-1	6	17	11	64	98
0	186	131	145	397	859
$+1$	25	32	20	72	149
$+2$	11	11	4	16	42
$+3$	2	6	4	10	22
$\geq +4$	7	11	4	14	36
MNE Total	242	222	193	602	1,259
$\bar{N}_{2000}$	1.49	2.38	1.56	1.96	
$\bar{N}_{1996}$	1.41	2.28	1.50	2.01	

Sources: MIDI universe 1996 and 2000 (not matched to USTAN). MNEs with regional presence of at least one affiliate in 1996; manufacturing MNEs and their majority-owned foreign manufacturing affiliates.
Notes: Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe). Median number of affiliates by MNE, location and year: 1.

location) in 1996 more than four keep exactly two locations (large-font entries in row 2). A similar pattern holds for any multiple-location MNE: entries along the diagonal exhibit the highest frequency in every row and every column. Regional expansions are gradual: the frequencies above the diagonal decrease monotonically in every row. Regional exits, however, are generally not gradual: MNEs that exit most frequently abandon all foreign locations at once; frequencies in the first column dominate frequencies below the diagonal in the third and fifth row (small-font entries in column 1). There is a remarkable number of complete withdrawals between 1996 and 2000 (477 out of 2,247 MNEs), but most of those withdrawers were present in only one foreign location in 1996 (365 out of 477). Note that the MIDI data cover the universe of German firms with FDI above minimum thresholds, and sample attrition is mitigated by the legal obligation to report and BuBa's commitment to follow up on missing questionnaires.

At the extensive margin, we query the number of affiliates and countries that are involved in changes to foreign presence. German MNEs typically pursue a single-affiliate strategy of foreign presence: the median number of affiliates per aggregate location is one. Table 3 shows that, once an MNE has established its presence in a given location with at least one affiliate, the number of affiliates hardly changes: 859 out of 1,259 MNE observations in given locations exhibit no change to the number of affiliates between 1996 and 2000; 247 out of 1,259 MNEs increase or decrease the number of affiliates by one. A small remainder of 153 parents chooses to change the number of affiliates by more. (The MNE total in Table 3 is smaller than that in Table 2 because we condition on presence in a location.) Changes to the number of host countries within locations are even less frequent than changes to the number of affiliates: an analysis of host country changes similar to Table 3 shows that 947 out of 1,259 observations of MNEs exhibit no change in the number of selected host countries within the aggregate location.[20] Motivated by these findings, we define the extensive margin as location selection in its most basic sense: an MNE's entry into an aggregate location with the first worker at its first affiliate.

B Wage data

We construct two types of wage variables by location: industry-median wages, and MNE-specific wages (wage premia or discounts relative to the industry medians). For industry-median wages in foreign locations, we use manufacturing wages by country and industry from the UNIDO Industrial Statistics Database 1996-2001 at the 3-digit ISIC level (dividing industry wage bills by employment) for our main analysis. We also report robustness checks using UBS wage data (UBS 2003) and OWW wage data (Occupational Wages around the World, Freeman and Oostendorp 2001). Appendix F.2 provides details on these wage data sources and the manufacturing occupations that we use. For industry-median wages in Germany, we take German wages from the same outside sources as affiliate wages to ensure comparability at the industry level.

We construct industry-median wages differently for the two margins to address econometric concerns. For intensive-margin labor-demand estimation on the second stage, we use median wages over industries by country. The median mitigates possible sectoral workforce composition effects behind local wages. Concretely, we take the arithmetic mean over the industry-median

[20]Infrequent net changes to the number of affiliates and countries could, in principle, conceal gross alternations such as changes to the country composition within a location or exit and reentry with a different affiliate. The data show that only small shares of MNEs that maintain a constant number of affiliates within a location change countries. In both CEE and WEU 4.2 percent of MNEs with constant affiliate numbers between 1996 and 2000 change host country, and 7.2 percent of the MNEs with constant affiliate numbers in DEV change country, but none do so in OIN. Similarly small fractions are associated with changing affiliate IDs, suggesting that the few gross alternations beyond net changes are mostly country changes and not reentries with different affiliates.

wages across the foreign countries where the MNE is present in a given year, and we take Germany-wide industry medians of the home wages by year. These wages are the decision-relevant local labor costs that the MNE faces at the intensive margin. For extensive-margin estimation on the first stage, wage variables must not depend on an MNE's country selection. Moreover, foreign wages are location-specific attributes and would therefore not be identified for the cross-section of MNEs in binomial choice models. To rely less on time variation, we make our foreign-wage variables (industry-median wages by country) MNE-specific relative to competitors: we take competitor averages for every MNE over the foreign wages that the MNE's German competitors pay.[21] The wage in CEE, for example, is the median wage averaged across the CEE countries where competitors' affiliates are located on the first stage. We apply the same procedure also to all other host-country characteristics. For the annual home wage, we use industry-mean wages because wage variables that reflect the workforce composition are valid predictors at the extensive margin.

For MNE-specific wages, we use USTAN data to obtain German parent wages and an outside data source for estimates of foreign affiliates' wage premia. The German parent's wage is the USTAN labor cost divided by German employment. MIDI does not report paid wages for foreign affiliates. To account for typical MNE wage premia on top of local labor costs abroad, we obtain Swedish affiliate wages by host country and industry from the IUI (Research Institute of Industrial Economics) data base for 1998 (Ekholm and Hesselman 2000), divide the MNE wages by the UNIDO manufacturing wages for host country and industry in 1998, and use the log ratios as our measures of wage premia at foreign locations. German and Swedish MNEs exhibit similar labor-demand behavior abroad (Becker et al. 2005).

C Estimation sample

Table 4 reports sample means over MNEs with presence in a given location. For CEE wages in second-stage labor-demand estimation, for instance, the table shows the location mean wage over the foreign countries where the MNEs are present. For CEE wages in first-stage selection estimation, the table shows log wages paid by the competitors of the MNEs with FDI in CEE.[22] In

[21] We consider only competitors within an MNE's broad manufacturing industry. The eight industries are: food; textiles and leather; wood, pulp and paper; chemicals, rubber, plastic and energy producing materials; mineral and metal products; machinery and equipment; transport equipment; and manufactures not elsewhere classified.

[22] We use the wage level at $t-\tau$ as a variable in selection estimation, not its log. For comparisons to the log wage at t in Table 4, we report the log of the sample-mean wage at $t-\tau$ ($\tau=2$).

Table 4: Sample Means of Variables

(t: 1998-2001, $t-\tau$: 1996-99)	HOM (1)	CEE (2)	DEV (3)	OIN (4)	WEU (5)
Indic.: Presence in t	1	.378	.323	.300	.702
Indic.: Presence in $t-\tau$	1	.351	.298	.283	.706
MNE-wide regressors (Labor-demand estimation)					
Wage bill share (t)	.791	.067	.050	.171	.192
ln Fixed assets (t)	17.267	14.893	15.112	15.804	15.281
ln Turnover (t)	18.449	15.936	16.511	17.281	17.071
ln Wage (t)	10.360	8.286	8.654	10.317	10.098
Competitor-average regressors (Selection estimation)					
ln sample-mean Wage ($t-\tau$)	10.428	8.278	8.708	10.348	10.076
Comp.s' hosts' ln Market access ($t-\tau$)	11.211	10.501	12.595	12.758	11.526
Comp.s' hosts' skill share < Home ($t-\tau$)	20.121	18.918	22.301	22.455	20.677
Comp.s' hosts' skill share ≥ Home ($t-\tau$)	41.988	38.962	47.854	49.371	43.271
Comp.s' hosts' distance ($t-\tau$)	31.606	29.445	35.811	36.369	32.548
Comp.s' hosts' ln Cons. p.c. ($t-\tau$)	30.389	28.559	33.904	34.373	31.183
Parent-firm regressors (Selection estimation)					
Indic.: Headquarters West Germany ($t-\tau$)	.973	.964	.974	.970	.975
ln Count of host countries ($t-\tau$)	1.138	1.331	1.637	1.475	1.263
Employment ($t-\tau$)	2,101	3,492	4,942	3,691	2,204
Fixed assets ($t-\tau$) [million]	239.3	451.6	637.1	499.7	273.1
Turnover ($t-\tau$) [million]	500.8	876.8	1,176.8	842.9	504.9
Intm. inputs ($t-\tau$) [million]	287.3	527.8	678.4	460.7	270.2
Liability ($t-\tau$) [million]	280.0	504.8	701.0	522.0	297.1
Parent observations	1,654	616	463	496	1,104

Sources: MIDI and USTAN 1996 to 2001, censored (second-stage) estimation sample of 1,654 MNEs.
Notes: Averages of MNE variables are conditional on presence. Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

our main estimation specification, we consider multinational labor demand during the years 1998-2001 for a sample of 1,654 MNEs and infer their location selection two years prior to production from an uncensored sample of 3,392 MNEs during the years 1996-1999.[23] For robustness checks, we will also use a single cross-section of 326 MNEs in 2000 and their location selection in 1996. The frequency of MNE presence abroad increases by two to four percentage points between 1996-99 and 1998-2001 in all locations except WEU (Western European countries), where it slightly falls in the censored panel.

German MNEs spend the bulk of their wage bill (79 percent) at home because German wages

[23] We lose observations on the second stage (t) mainly because of missing wage information at affiliate locations, whereas competitor-mean wages on the first stage ($t-\tau$) are less sensitive to missing information.

and German employment are relatively high compared to foreign locations. From German MNEs, CEE receives labor expenditures beyond the remaining developing world combined. (Note that shares do not add to unity across columns because averages are conditional on presence, omitting absent MNEs). A similar cross-location pattern arises for turnover and capital stocks. Substantial wage disparities persist across locations. Between Germany and CEE, for instance, MNE wages differ by 2.1 log points, or a factor of around 800 percent ($\exp\{10.360 - 8.286\} = 8.0$ for 1998-2001). This MNE-level difference is smaller, however, than the country-population weighted wage gap of about 1,000 percent ($1/.099$) in the raw UNIDO wage data in 2000 (the population-weighted wage gap in OWW data is almost the same with $1/.098$). The smaller conditional differential is consistent with MNE selection into relative high-wage countries within the low-wage region CEE (Marin 2004).

German MNEs in CEE, compared to any other location, face competitors in host countries that offer the least market access, that have the smallest skill endowments, that are geographically the closest and that exhibit the smallest per-capita consumption. The CEE wages paid by German competitors of MNEs in CEE are below those paid by German competitors in DEV. MNEs in OIN, at the other extreme, face German competitors with the strongest host-country market access and host-country skill endowments.

Parent-level covariates are suggestive of selectivity effects at their means. Parents with headquarters in East Germany (including Berlin) are slightly more likely to expand to CEE and OIN than the average German MNE. Conditional on their presence abroad, MNEs exhibit larger home workforces, larger parent-firm equity or debt, and higher parent-firm capital-labor ratios.

III Estimation

We turn to estimation of the MNE model. The model is a system of foreign location choices at the extensive margin and of labor-demand outcomes by location at the intensive margin. The model permits estimation of well defined extensive-margin elasticities, and naturally corrects intensive-margin estimates for extensive-margin adjustment. We estimate the model parametrically and non-parametrically. We control for all relevant location alternatives and condition on a comprehensive set of cost-function variables as well as MNE wage premia over the local median wage. Exclusion restrictions and timing provide identification.

A Identification

The labor-demand outcome on the second stage is separately identified from location selection on the first stage because the MNE chooses current output, employment and capital in response to news after location choice, whereas location-selection estimation is based on past information and a separate set of parent-firm and competitor-level variables. Location selection on the first stage is separately identified from labor demand on the second stage because parent-firm variables and competitor-level host-country attributes at decision time are among the predictors of future presence but not directly relevant for operation on the second stage other than through the propensity of presence. Output is a regressor in cost function estimation, so no identifying assumptions on output responses under product-market competition are needed.

We use industry-wide wages to identify employment responses at the extensive and the intensive margin. To control for MNE-specific wage premia beyond industry-wide wages at the intensive margin, we include the MNE wage residual that is orthogonal to presence propensities. Industry-wide German home wages are plausibly exogenous to the individual MNE at both margins because German manufacturing firms face bargained wage schedules from industry-specific collective agreements between employer associations and unions. The threat of employment relocation abroad potentially affects the outcome of collective wage bargaining. We control for the employers' propensity to select into foreign locations in parametric and nonparametric two-stage approaches so that coefficients on industry-wide German home wages are adequately identified at both margins from cross-sectoral variation. Time variation in home wages provides additional identification. Industry-wide foreign labor costs are wage medians by location and also exogenous to the MNE. Foreign affiliates of German MNEs are few and small, and observed wage premia at MNEs over their local competitors do not support canonical monopsony models of market power abroad. For selection estimation on the first stage, competitors' median labor costs by location vary across MNEs by construction, and time variation provides additional variation. For labor-demand estimation on the second stage, median foreign wages provide identification in the MNE cross section because MNEs' country choices within aggregate locations differ so that the exposure to median foreign wages varies across firms and over time.

Serial correlation in the selection disturbance, due to persistence in unobserved local market conditions say, could contribute to the observed hysteresis of foreign presence. We therefore perform estimation under varying assumptions on serial correlation and consider different time horizons of location selection. Repeated MNE cross sections with two-year selection-outcome lags are our benchmark. We also obtain results under a second-order autoregressive error component in location selection, as well as other autocorrelation specifications, and obtain results for a single cross-section of firms with location selection at a four-year lag. An Akaike information criterion indicates that independent errors receive most empirical support.[24]

Unobserved MNE heterogeneity is a concern but mitigated in our framework and data. Our labor-demand estimation controls for MNE-specific wage residuals and, consistent with cost-function estimation, conditions on output. So we explicitly account for the heterogeneity in product-market shares. We use current capital-stock observations as regressors in empirical analysis, viewing capital as pre-determined during location selection, and so control for differences in capital use. We include a large set of time-varying parent-level variables in selection estimation on the first stage—among them MNE size and financial measures. On the second stage, inverse Mills ratios or nonparametric propensities of foreign presence control for heterogeneity and the MNEs' motives to conduct FDI. In addition to their manufacturing affiliates abroad, MNEs may operate commercial affiliates whose MNE-specific presence could affect results. We have repeated our analysis for the full sample of foreign affiliates in any sector and find similar sign patterns for cross-wage elasticities, indicating employment substitutability when significant. A remaining unobserved MNE-specific performance advantage, if not shared with workers through residual wage payments, would arguably cause domestic and foreign employment to expand simultaneously and suggest a bias of labor-demand elasticity estimates towards complementarity. But we consistently find cross-regional substitutability.

[24]Closely related to autocorrelation is the consideration of potential adjustment costs, as MNEs expand across countries within aggregate locations or as MNEs open additional plants within countries. These adjustments go beyond our basic extensive margin of location selection with the first employee at the first affiliate and are akin to a decomposition of the current intensive margin into additional extensive margins. In augmented regressions that condition on lagged employment (reported in Subsection E) or future output (not reported for brevity), we find broadly similar cross-wage elasticity estimates at both margins and infer that the existence of additional extensive margins behind our current intensive margins does not seem to affect our estimate of the basic extensive margin.

B Location choice

We first estimate location-selection equations (6)

$$d^{\ell}_{jt} = \mathbf{1}\left(H(\mathbf{z}_{j,t-\tau}) + \eta^{\ell}_{j,t-\tau} > 0\right).$$

Probit estimation. We start by investigating the implication of Assumption A that the selection-shock covariances between locations are constant for all locations. We obtain estimates for the covariances from multivariate probit estimation of simultaneous selection into the four foreign locations (on the same set of regressors as in Table 5). We fail to reject joint equality of the six correlation coefficients between the four equations with a χ^2 test statistic of 4.63 (p value .592).

The plausibility of Assumption A verified, we turn to probit estimation location by location and investigate alternative specifications for serial correlation. To pick the serial-correlation specification with most empirical support, we apply the Pan (2001) extension of Akaike's information criterion to the general-estimation-equations probit quasi-likelihood function. By this measure, time independence of the disturbances receives most support in every single location (compared to hypothesized AR(1), AR(2) and stationary processes of the disturbances). We therefore choose ordinary probit estimates as our benchmark, but will also report labor-demand elasticities under the alternative assumption of AR(2) disturbances in Section E.[25]

Table 5 presents the probit results as marginal effects. Among the firm-level predictors, we include interactions between past presence indicators and wages to capture a potentially different effect of the wage differential on an MNE with presence at a location. Past presence elsewhere (off the diagonal) has little predictive power, but past presence for the location itself is typically a statistically significant and salient predictor of presence (excepting OIN where the wage-presence interaction takes over). Indicators of past presence also control for permanent but unobserved MNE characteristics, such as lasting productivity or ownership advantages. When leaving interactions between wages and past presence out for a comparison, past presence at the same location has a highly statistically significant probability effect of .779 (standard error .022) in CEE, .671 (.027) in DEV, .713 (.026) in OIN, and .747 (.020) in WEU. The importance of past presence at the two-year horizon is consistent with sunk costs and hysteresis in location choice.

[25] An AR(2) specification ranks second or third in terms of the information criterion, depending on location, compared to independence, to an AR(1) and to a stationary two-year lag specification.

Table 5: MARGINAL EFFECTS IN POOLED PROBIT REGRESSIONS

Predictors $(t-2)$ Presence (t)	CEE (1)	DEV (2)	OIN (3)	WEU (4)
FDI in CEE $(t-\tau)$	.609	.222	.430	-.388
	(.234)***	(.275)	(.298)	(.287)
FDI in DEV $(t-\tau)$	.015	.740	-.099	-.093
	(.110)	(.129)***	(.072)	(.150)
FDI in OIN $(t-\tau)$	-.307	-.571	-.067	-.076
	(.413)	(.323)*	(.478)	(1.046)
FDI in WEU $(t-\tau)$	.309	.133	.087	.987
	(.202)	(.287)	(.252)	(.016)***
Home sector wage	.0008	.003	.007	.013
	(.004)	(.004)	(.003)**	(.007)*
FDI in loc. $(t-\tau)$ × Home sector wage	.002	-.003	-.015	-.015
	(.005)	(.004)	(.004)***	(.007)**
Competitors' wages CEE	-.053	-.016	.002	-.094
	(.054)	(.045)	(.040)	(.058)
FDI in CEE $(t-\tau)$ × Comp.s' wages CEE	.035	-.068	-.087	.099
	(.065)	(.057)	(.051)*	(.082)
Competitors' wages OIN	-.004	.000006	-.026	.032
	(.014)	(.016)	(.015)*	(.020)
FDI in OIN $(t-\tau)$ × Comp.s' wages OIN	.013	.036	.035	.001
	(.027)	(.026)	(.019)*	(.033)
ln Count of host countries	.068	.131	.057	.158
	(.039)*	(.035)***	(.028)**	(.054)***
Employment $(t-\tau)$ [thsd]	.019	.022	.005	-.017
	(.009)**	(.008)***	(.006)	(.017)
Turnover $(t-\tau)$ [billion]	-.012	.016	.057	.933
	(.064)	(.051)	(.029)*	(.230)***
Intm. inputs $(t-\tau)$ [billion]	.016	-.064	-.085	-1.086
	(.073)	(.059)	(.037)**	(.272)***
Liability $(t-\tau)$ [billion]	-.173	-.073	-.006	-.362
	(.073)**	(.071)	(.053)	(.122)***
Obs.	2,459	2,459	2,459	2,459
Pseudo R^2	.551	.519	.546	.452

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages), pooled sample of manufacturing MNEs and their majority-owned foreign manufacturing affiliates with two-year selection lags ($\tau = 2$).
Notes: Standard errors in parentheses: * significance at ten, ** five, *** one percent. Further regressors (not significantly different from zero at five percent level in any location): Competitors' wages DEV and WEU and their interactions with FDI presence in DEV and WEU, Competitors' hosts ln Market access, Indic. Headquarters West Germany, Fixed assets, Competitors' hosts skill share < Home, Competitors' hosts skill share ≥ Home, Competitors' hosts distance, Competitors' hosts ln Cons. per capita. Without wage-presence interactions, past presence has a marginal effect of .779 (standard error .022) in CEE, .671 (.027) in DEV, .713 (.026) in OIN, and .747 (.020) in WEU. Locations: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

The home wage has the expected positive sign in all regressions and is a statistically significant predictor for presence in OIN and WEU, both by itself and in its interaction with past presence. The negative coefficients on the interaction terms suggest that wage differentials matter less for the location decision of MNEs that already own an affiliate in the region. But positive net effects prevail at sample mean presence frequencies (which are .35 in CEE, .3 in DEV and OIN, and .7 in WEU by Table 4). With home wages already controlling for the foreign-to-home wage differential, several foreign wages are statistically insignificant predictors. Insignificant coefficients of foreign wages are common in the literature on location choice (e.g. Devereux and Griffith (1998) for U.S., and Buch et al. (2005) for German MNEs). We need only home-wage coefficients for the cross-wage elasticities at the extensive margin (there is no extensive margin for home where foreign wages would enter). Ultimately, even the statistically weak home-wage prediction for CEE will turn out to contribute to a statistically significant cross-wage elasticity in bootstraps. Recall that the wage impact at the extensive margin is the product of two factors, the home wage effect on foreign presence (from Table 5) times the effect of foreign presence on foreign employment (to be estimated).

We include a large set of MNE and host-country variables. MNE characteristics are statistically highly significant predictors of location choice with p-values on the χ^2 statistics of .05 or below, except for OIN. German MNEs with large home employment, high turnover, low intermediate-input levels and low parent debt are more likely to be present at most or all foreign locations. The MNE's number of host countries in the past significantly raises the likelihood of presence. An indicator of parents' headquarters in West Germany and parent fixed assets, however, are not statistically significant at the five percent level in any location.

For location-specific variables, χ^2 tests exhibit a mixed pattern with p values between .04 (DEV) and .75 (CEE). Table 5 does not report the covariates for brevity. The suppressed regressors include wages in DEV and WEU and their interactions with past presence in DEV and WEU, host-market access, host-country skill shares, host-country distance, and host-country per-capita consumption. Although we transform all location covariates to the competitor level for the relevant cross-sectional variation MNE by MNE, none of the location-specific covariates is individually significant at the five percent level in any location after conditioning on location wages.[26]

[26]To tentatively control for an outside margin of arm's length trade between independent firms, we also included a set of sector and location specific import and export measures but found the trade variables not to be statistically

We run the same regression as in Table 5 on the single cross section of MNEs in 2000 for a four-year horizon, using selection predictors in 1996. Results are broadly similar and suggest that cross-sectional variation in wages drives our results. We report the according labor-demand cross elasticities in Section E.

Nonparametric propensity score estimation. To break the curse of dimensionality, we choose seven core predictors and a polynomial approximation around them, while we linearly condition on the set of remaining firm and host-country variables. For the choice of the seven core variables we use existing evidence in the FDI literature to guide us: market access (Head and Mayer 2004) and the count of an MNE's past host countries (Buch et al. 2005) are regarded as important predictors. For purposes of our estimation, wages in the five aggregate locations belong among the core variables. To query the appropriate order of the polynomial expansion around the core variables, we use two criteria. Cross validation lends slightly more support to a second-order polynomial in the core variables. But F tests show that more wage predictors are statistically significant in a third-order polynomial specification. We report nonparametric results from a third-order polynomial expansion here, yet ultimate elasticity estimates differ little. As to serial correlation, we find the specification with independent disturbances to exhibit a better fit than serial correlation, similar to probit estimation.

Table 6 reports coefficient estimates by location. The predicted propensity scores of location choice are .334 for CEE, .288 for DEV, .261 for OIN and .612 for WEU—slightly under-predicting the actual frequencies of presence in Table 4 but reflecting the relative frequencies across locations. Marginal effects are close to those in the probit regressions. Estimates of past presence indicators along the diagonal continue to have a magnitude similar to probit estimation. When leaving interactions between wages and past presence out, past presence at the same location has a highly statistically significant probability effect of .759 (standard error .018) in CEE, .668 (.020) in DEV, .711 (.017) in OIN, and .707 (.024) in WEU. Inclusion of wage interactions with past presence shifts much predictive power to the interaction terms in DEV and all predictive power to the interaction terms in OIN. In WEU, the interaction term countervails the high marginal effects of past presence.

significant predictors of location choice; we leave them out of the regressions in Table 5. Results are robust to the inclusion of year dummies.

Table 6: MARGINAL EFFECTS IN NONPARAMETRIC PROBABILITY MODEL

Presence (t) Predictors ($t-2$)	CEE (1)	DEV (2)	OIN (3)	WEU (4)
FDI in CEE ($t-\tau$)	.634	.110	.201	-.158
	(.144)***	(.148)	(.138)	(.184)
FDI in DEV ($t-\tau$)	-.047	.340	-.079	-.010
	(.087)	(.115)***	(.083)	(.107)
FDI in OIN ($t-\tau$)	.022	.042	.054	.281
	(.551)	(.564)	(.551)	(.685)
FDI in WEU ($t-\tau$)	.186	-.033	-.033	1.229
	(.221)	(.215)	(.203)	(.259)***
Series terms involving wages: p-values from F tests				
Home sector wage terms		.030	.007	.094
Competitors' CEE wage terms				
Competitors' DEV wage terms				
Competitors' OIN wage terms	.005	.103		
Competitors' WEU wage terms	.056			
Employment ($t-\tau$) [thsd]	.014	.011	-.009	-.015
	(.006)**	(.006)*	(.006)	(.008)**
Turnover ($t-\tau$) [billion]	.004	.078	.251	.415
	(.061)	(.062)	(.059)***	(.075)***
Intm. inputs ($t-\tau$) [billion]	-.003	-.140	-.303	-.444
	(.068)	(.070)**	(.066)***	(.085)***
Liability ($t-\tau$) [billion]	-.137	-.026	.004	-.179
	(.046)***	(.047)	(.044)	(.056)***
Competitors' hosts ln Cons. p.c. ($t-\tau$)	.079	.013	-.010	.023
	(.030)***	(.031)	(.029)	(.037)
Obs.	2,459	2,459	2,459	2,459
R^2	.662	.617	.630	.553

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages), pooled sample of manufacturing MNEs and their majority-owned foreign manufacturing affiliates with two-year selection lags ($\tau = 2$).
Notes: Standard errors in parentheses: * significance at ten, ** five, *** one percent. Third-order polynomials in Wages, $\ln$ Count of host countries, Competitors' hosts' $\ln$ Market access. Further regressors (not significantly different from zero at five percent level in any location): Interactions of competitors' wages with FDI presence, $\ln$ Host count, Competitors' hosts $\ln$ Market access, Indic. Headquarters West Germany, Competitors' hosts skill share, Competitors' hosts distance. Without wage-presence interactions, past presence has a marginal effect of .759 (standard error .018) in CEE, .668 (.020) in DEV, .711 (.017) in OIN, and .707 (.024) in WEU. Locations: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

Table 6 presents F-tests of joint significance of individual wages for p values at or below the .1 threshold. Similar to probit estimation, polynomial terms that involve home wages predict location choice more successfully than most foreign wages (except OIN wages). Home wages are the predictors we need for cross elasticities at the extensive margin. Series terms involving the home sector wage predict selection into DEV and OIN at the five percent significance level. Significant parent-level covariates from probit estimation remain significant predictors under nonparametric estimation, excepting the host country count variable. Similarly, statistically insignificant parent-level covariates remain insignificant, and insignificant host-country variables continue insignificant.

C Labor demand estimation with selectivity correction

We proceed to estimate employment outcomes (2)

$$y_{jt}^{\ell} = \mathbf{x}_{jt}^{\ell}\beta^{\ell} + m^{\ell}(\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau})) + \boldsymbol{\psi}_{jt}\theta^{\ell} + \epsilon_{jt}^{\ell}$$

for all locations. $\mathbf{x}_{jt}^{\ell}$ is the vector of observations-specific labor-demand predictors from eq. (1), $\boldsymbol{\psi}_{jt}$ is the vector of the MNE's wage residuals and orthogonal to presence predictors $\mathbf{z}_{j,t-\tau}$ by construction (Appendix E). We stack MNE observations with different presence choices abroad by setting regressors for locations of absence to zero, and include absence indicators accordingly (see Appendix B for natural assumptions underlying stacking). $m^{\ell}(\cdot)$ is the relevant presence propensity predicted by $\mathbf{z}_{j,t-\tau}$ (inverse of the Mills ratio under Assumption A, predicted propensity scores under Assumption B). We use cross validation to choose the order of polynomial expansion at this stage. A third-order approximation performs better in the more remote locations DEV and OIN (but worse in CEE and WEU); we use a third-order approximation because the extensive margin is potentially more relevant at remote locations. In both parametric and nonparametric regressions we include absence indicators among the predictors to prevent stacking bias.

We implement the second-stage estimation for all but one location (excluding home) by iterating Zellner's (1962) seemingly unrelated regression (SUR) over the estimated disturbance covariance matrix until the estimates converge. This is equivalent to maximum-likelihood estimation (Dhrymes 1971) and makes estimation invariant to the deleted location equation (Barten 1969). Through constraints, we impose linear homogeneity in factor prices and symmetry of wage coefficients (see Appendix A), and recover coefficients of the deleted home equation. We treat induced heteroskedasticity following Heckman (1979), resulting in different standard errors on symmetric coefficients.

We obtain estimates of translog cost function equations for 1,654 stacked MNE observations between 1998 and 2001. The estimation equations include the full sets of wage, turnover and fixed asset regressors, the scaled equivalent of the constant, and indicators of absence from all other locations. All but two wage coefficients are significantly different from zero at the one percent level, and all wage coefficients but one are significant at the five percent level in parametric and nonparametric regression (see Table G.1 in the Appendix). Most coefficients on turnover and fixed assets are similarly highly significant.

Under Assumption A, the predicted selectivity hazards (inverses of Mills ratios) are statistically different from zero at the five or ten percent level in all equations except CEE. Recall that what ultimately matters for sign and significance of cross-elasticities at the extensive margin is the product between the home wage effect on foreign presence (from selection estimation before) and the effect of foreign presence on foreign employment (the selectivity hazard coefficient in the labor-demand regression here times the first derivative of the inverse of the Mills ratio). As will become clear shortly, individually small factors do not adversely affect statistical precision of the overall product behind the extensive margin. The cross-elasticity estimates at the intensive margin will be statistically different from the total elasticity for all locations in our data, corroborating the importance of the extensive margin. Under Assumption B, we use third-order polynomials in the predicted propensity scores for all locations. χ^2 tests on the series terms overwhelmingly reject their joint equality to zero. The translog cost function regressors predict the bulk of labor demand variation across locations, with the R^2 goodness of fit ranging between .95 and .99 across equations under parametric and nonparametric selectivity correction.

Elasticities of multinational labor substitution. Table 7 shows own-wage and cross-wage substitution elasticities for permanent wage changes by one percent in different locations, separately for the extensive and the intensive margins. There is no well-defined extensive margin for selection into the home location in a sample of MNEs that are observed only if active in the home location. One margin at a time is set to zero to isolate the effect at the other margin. While the plain log wage effects on wage bill shares are additive over the two margins, cross-wage substitution elas-

Table 7: CROSS-WAGE ELASTICITIES UNDER PARAMETRIC SELECTIVITY

Employment change (%) in		Wage change (by 1%) in HOM (1)	CEE (2)	DEV (3)	OIN (4)	WEU (5)
HOM	*intensive*	-.307**	.026***	-.003	.085	.198***
CEE	*intensive only*	.820***	-.932***	-.288***	.365***	.035
	extensive only	.794***	-1.029***	.021	.041	.084
DEV	*intensive only*	-.157	-.514***	-.179	.679***	.171
	extensive only	.857***	-.149	-.988***	.362	.437
OIN	*intensive only*	1.303	.179***	.186***	-2.630**	.961***
	extensive only	.629***	.169	.009	-.157	.052
WEU	*intensive only*	1.205***	.007	.019	.383***	-1.614***
	extensive only	.838***	-.098	.057*	.574	-.880***

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages).
Notes: Elasticities at the extensive and intensive margins from 1,654 stacked MNE observations. Underlying labor demand estimates from parametric selectivity-corrected ISUR estimates (Assumption A, Tables 5 and G.1). Standard errors from 200 bootstraps: * significance at ten, ** five, *** one percent. Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

ticities are not additive by eq. (4). We bootstrap 200 times over joint selection (6) and outcome (2) estimation to infer elasticities of labor substitution (4) and their standard errors at both margins.[27]

Own-wage elasticities along the diagonal—for both intensive and extensive margins—are uniformly negative, as production theory requires. While this might be expected for estimates at the intensive margin, it is a reassuring finding for estimates at the extensive margin. As is common, we impose linear homogeneity in factor prices and symmetry of wage coefficients at the intensive margin through constraints on the translog regression. But we do not restrict estimates at the extensive margin—neither under parametric nor nonparametric selectivity correction. The own-wage elasticity of substitution at home of $-.31$ falls into the established range in the labor-demand literature. The own-wage elasticity of substitution in foreign locations is considerably larger than at home, suggesting that MNE employment abroad responds more sensitively to labor costs there than home employment responds to home wages. We are not aware of results in the literature to which we could compare the own-wage elasticities in foreign locations along the diagonal.

Cross-wage elasticities in the first row (foreign wage effects on home employment) and in the

[27] Bootstrapping is advantageous because it does not require treatment of insignificant wage coefficients from the first stage to quantify the extensive margin. Moreover, Eakin, McMillen and Buono (1990) show in simulations that analytic confidence intervals for elasticity estimates can differ considerably from bootstrapped confidence intervals.

first column (home wage effects on foreign employment) are significantly positive for eight out of twelve estimates at the intensive and extensive margins. A one-percent reduction in the wage in CEE, for instance, is associated with a .03 percent drop in home employment at German MNE parents. In contrast, a one-percent increase in the German sector wage is associated with a .82 percent boost to MNE employment in CEE at the intensive margin and a .79 percent boost at the extensive margin. So, home and CEE employment are substitutes within MNEs. The large difference in cross-wage effects between first row and first column is consistent with two stylized facts. First, employment at German MNE parents is larger in levels than at their CEE affiliates so that a smaller percentage wage drop in Germany means a larger reduction in employment abroad in absolute terms. Second, CEE workers have an arguably lower labor productivity than German workers so that CEE employment levels are more responsive to a given foreign wage change.

The extensive margin is important in every location. Extensive-margin elasticities in the first column are strictly positive and statistically different from zero everywhere. So, home and foreign employment are substitutes within MNEs not only at the intensive but also at the extensive margin. The CEE and DEV home wage effects on selection were not statistically different from zero on the first stage with probit (Table 5). But when combined with the selection effect on labor demand on the second stage (selectivity hazard coefficient in Table G.1 times the first derivative of the inverse of the Mills ratio), all four extensive-margin elasticities become statistically different from the respective total elasticities. Recall that the total cross-wage elasticity in eq. (4) is the sum of the marginal wage-bill changes at the two margins and the home wage bill share. We reject equality between the intensive-margin elasticity alone (setting the extensive-margin response to zero) and the total elasticity for every location (with t statistics between 2.6 and 8.4). In summary, despite their small and hard-to-detect magnitudes, extensive-margin effects are statistically important. That is not the case for all intensive-margin elasticities.

Intensive-margin cross-wage elasticities in the first row and column are not significantly different from zero in four out of eight cases, and those are exactly the intensive-margin elasticities for the distant locations DEV and OIN. For those distant locations, only the extensive margin shows statistically significant employment responses to permanent wage changes.

Cross-wage estimates beyond the first row and column are statistically different from zero in some cases. Notable pairs of mutually positive cross-wage effects at the intensive margin are OIN on WEU (.38) and vice versa (.96), as well as DEV on OIN (.19) and vice versa (.68). A positive cross-wage effect indicates substitutability of employment between OIN-WEU and OIN-DEV. A striking pair of negative cross-wage effects at the intensive margin is CEE on DEV (-.51) and vice versa (-.29), presenting the single case of statistically significant employment complementarity in our data. Assuming that MNEs with joint presence in CEE and DEV operate vertical production chains, employment complementarity might be explained with mutually provided producer services and intermediate inputs as in Markusen (1989). Cross-wage effects at the extensive margin are not statistically significant beyond the first row and column. Together, these findings suggest that a main motive for MNE formation is labor substitution with the home location at the extensive margin but that subsequent MNE operation draws on both employment substitutability and complementarity across established global locations at the intensive margin.

D Comparison to earlier approaches

A core aspect of our integrated approach is the estimation of a global labor-demand system. Compared to single-equation home labor demand estimation in the earlier literature, this integrated approach leads to substantively different results even in the absence of extensive-margin correction. Table 8 compares intensive-margin specifications, moving from a parsimonious specification all the way to system estimation without extensive-margin correction. Column 1 displays a typical home labor demand equation for European MNEs as in Braconier and Ekholm (2000), Konings and Murphy (2006), or Marin (2004). The specification includes wages in affiliates in other European regions and conditions on a single output measure (home turnover here, value added in some earlier papers). As do Konings and Murphy (2006), we also condition on a full set of MNE presence indicators in this and all other columns. Wages in CEE and WEU are not statistically significant predictors, and have a negative sign. In column 2 we additionally control for MNE turnover in other world regions. This is similar to papers controlling for worldwide output, but differs in controlling separately for output from different world locations. Wages in CEE and DEV now become positive, pointing to substitutability, but among the foreign wages only the WEU wage turns statistically significant. Beyond Braconier and Ekholm (2000), Konings and Murphy (2006) and Marin (2004), we control in column 3 for the MNE's capital stocks at home and in the four foreign locations as inverse proxies to unobserved user costs of capital. Point estimates on

Table 8: SINGLE- AND MULTIPLE-EQUATION ESTIMATION OF THE INTENSIVE MARGIN

HOM employment intensive margin	SINGLE HOM EQUATION (1)	(2)	(3)	(4)	SYSTEM (5)
ln *Wages*					
HOM	-.950	-.993	-1.011	-.991	-.333
	(.018)***	(.022)***	(.022)***	(.024)***	(.071)***
CEE	-.021	.024	.042	.043	.030
	(.020)	(.033)	(.033)	(.033)	(.004)***
DEV				.072	.005
				(.024)***	(.006)
OIN				-.033	.093
				(.038)	(.049)*
WEU	-.009	.047	.075	.092	.204
	(.007)	(.020)**	(.021)***	(.022)***	(.035)***
ln *Turnover*					
HOM	yes				
HOM-WEU		yes	yes	yes	yes
ln *Capital*					
HOM-WEU			yes	yes	yes
Obs.	2,141	2,141	2,141	2,141	2,141

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages).
Notes: Standard errors in parentheses: * significance at ten, ** five, *** one percent. All specifications include current presence or absence indicators (referred to as MNE-location FE by Konings and Murphy 2006). Not reported: Turnover, Capital Stocks, Current presence indicators, and Constant. Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

home and foreign wages become larger and gain in statistical significance. In column 4, we add DEV and OIN wages to take into account labor costs in relevant foreign locations outside Europe. The coefficient on the WEU wage turns even larger than before, whereas the own-wage coefficient for the home location remains close to negative one—a high magnitude compared to conventional labor demand estimates between -.3 and -.6.

The final column 5 displays estimates from a global translog labor-demand system, using cross-equation restrictions consistent with theory and an equivalent to maximum-likelihood estimation. The own-wage estimate drops by two thirds to a more commonly observed magnitude of -.3, whereas the WEU wage coefficient more than doubles, the CEE wage becomes statistically significant, and the DEV wage insignificant. So system estimation, which controls for all relevant locations and production factors, alters coefficient estimates considerably and is a prime reason for

Table 9: RELATIVE DIFFERENCE BETWEEN INTENSIVE-MARGIN ESTIMATES

Relative difference in employment effect estimates		Wage change in HOM	CEE	DEV	OIN	WEU
		(1)	(2)	(3)	(4)	(5)
HOM	*intensive*	.084†	.146†	-2.692	.093	.032†
CEE	*intensive*	.047†	.122†	-.566†	-.286†	.465
DEV	*intensive*	-2.728	-.515†	2.777	-.243†	-.281
OIN	*intensive*	.175	-.160†	-.204†	.126†	.178†
WEU	*intensive*	.092†	.697	-.255	.160†	.107†

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages), manufacturing MNEs and their majority-owned foreign affiliates in manufacturing.
Notes: The relative difference is the difference between the uncorrected and the selectivity-corrected (Assumption A) elasticity estimate, divided by the selectivity-corrected estimate. There are 2,141 stacked MNE observations for uncorrected ISUR and 1,654 for selectivity-corrected ISUR estimation. Daggers indicate cases where both the uncorrected and the selectivity-corrected estimate are statistically significant at the 5-percent level. Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

differences between our estimates and those in the earlier literature.

With the adoption of a labor-demand system, rigorous extensive-margin estimation of cross-wage elasticities becomes possible. Table 9 shows for the system of cross-wage elasticities how intensive-margin estimates change when statistically corrected for employment responses at the extensive margin. Every entry reports the relative change: the conventional estimate less the selectivity-corrected estimate, divided by the selectivity-corrected estimate. Uncorrected cross-wage elasticities are distorted towards complementarity in ten out of 25 cases in our data (negative signs) and, in two instances out of these ten, estimates are reversed from substitutability into outright complementarity (relative differences of less than negative one). Among the 16 coefficients that are statistically significant before and after extensive-margin correction (marked with daggers), not correcting for the extensive margin would result in under-estimated coefficients for six cross-wage elasticities—a bias towards complementarity in those cases and a bias towards substitutability in the remaining cases.

Following Harrison and McMillan (2006), we also split the sample into MNEs in industries with no significant intra-firm trade (horizontal FDI) and with significant intra-firm trade (vertical FDI); we do not find home employment to be a statistically significant complement to foreign employment in any industry or location, whereas Harrison and McMillan (2006) find complemen-

Table 10: FOREIGN-WAGE ELASTICITIES OF HOME EMPLOYMENT

Home employment change (%)	Wage change (1%) in HOM	CEE	DEV	OIN	WEU	Obs.
	(1)	(2)	(3)	(4)	(5)	(6)
Stacking						
Ass. A, UNIDO 98-01	-.307	.026	-.003	.085	.198	1,654
	(.131)**	(.005)***	(.008)	(.076)	(.063)***	
Ass. A, UNIDO 00	-.537	.029	.009	.301	.198	326
	(.252)**	(.018)	(.017)	(.188)	(.095)**	
Ass. A AR(2), UNIDO 98-01	-.300	.026	-.003	.084	.194	1,654
	(.198)	(.009)***	(.008)	(.112)	(.091)**	
Ass. A, UNIDO 98-01, lag y	-.307	.027	-.005	.111	.175	1,654
	(.112)***	(.006)***	(.008)	(.073)	(.054)***	
Ass. A, UBS 98-01	-.260	.014	.0009	.062	.183	1,628
	(.125)**	(.004)***	(.013)	(.081)	(.056)***	
Ass. A, OWW 98-01	-.303	.036	.010	.163	.094	1,467
	(.119)**	(.008)***	(.003)***	(.081)**	(.047)**	
Ass. B, UNIDO 98-01	-.317	.027	.004	.081	.204	1,654
	(.096)***	(.005)***	(.008)	(.065)	(.041)***	
Omnipresent MNEs						
Ass. A, UNIDO 98-01	-.152	.002	.059	.090	.0003	96
	(.376)	(.028)	(.055)	(.185)	(.222)	

Sources: MIDI and USTAN 1996 to 2001 (UNIDO, UBS and OWW wages).
Notes: Elasticities of wage effects on home employment (first row of elasticity matrix) at the intensive margin. Standard errors from 200 bootstraps: * significance at ten, ** five, *** one percent. Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

tarity for wages in low-income locations and home employment. Harrison and McMillan (2006) restrict the sample to manufacturing affiliates, as we do. The different findings for vertical FDI industries and affiliates in low-income locations may be due to differences in economic behavior between U.S. MNEs and German MNEs, or due to empirical method.

E Alternative specifications

To assess the robustness of our estimates, we compare several specifications and report the first rows of the cross-wage elasticity matrices (foreign wage effects on home employment) in Table 10, and the first columns separately by intensive and extensive margin in Tables 11 and 12 (home wage effects on foreign employment).

Table 11: HOME-WAGE ELASTICITIES AT THE INTENSIVE MARGIN

	Home wage change (1%), by regression specification							
	Stacking							Omnipr.
	UNIDO 98-01	UNIDO 00	UNIDO 98-01 AR(2)	UNIDO 98-01 lag y	UBS 98-01	OWW 98-01	UNIDO 98-01	UNIDO 98-01
Emplmt. chg. (%)	Ass. A	Ass. A	Ass. A	Ass. A	Ass. A	Ass. A	Ass. B	Ass. A
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
HOM	-.307	-.537	-.300	-.307	-.260	-.303	-.317	-.152
	(.131)**	(.252)**	(.198)	(.112)***	(.125)**	(.119)**	(.096)***	(.376)
CEE	.820	.834	.796	.842	.683	1.058	.834	.084
	(.157)***	(.528)	(.251)***	(.192)***	(.177)***	(.218)***	(.158)***	(1.056)
DEV	-.157	.400	-.146	-.295	.034	.959	.245	.978
	(.468)	(.793)	(.491)	(.473)	(.489)	(.252)***	(.428)	(.963)
OIN	1.303	3.811	1.280	1.696	.770	2.716	1.240	.320
	(1.183)	(2.420)	(1.812)	(1.080)	(1.002)	(1.392)*	(.990)	(.661)
WEU	1.205	1.117	1.178	1.063	.988	.889	1.244	.001
	(.382)***	(.529)**	(.551)**	(.325)***	(.299)***	(.438)**	(.241)***	(.826)
Obs.	1,654	326	1,654	1,654	1,628	1,467	1,654	96

Sources: MIDI and USTAN 1996 to 2001 (UNIDO, UBS and OWW wages).
Notes: Elasticities of home wage effects on foreign employment (first column of elasticity matrix) at the intensive margin. Standard errors from 200 bootstraps: * significance at ten, ** five, *** one percent. Locations: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

Foreign-wage elasticities of home employment in Table 10 are robust across specifications. Estimates on our benchmark sample (first row) with UNIDO wages and MNEs between 1998 and 2001 under Assumption A conform closely to several other specifications. The similarity between the 1998-2001 MNE sample and the single cross section of MNEs in 2000 (with location choice in 1996) in the second row is consistent with the view that cross sectional and not time series variation is the main source of identification at the intensive margin. The third row shows that an AR(2) error specification in the selection equation results in only minimal differences in cross-wage elasticity estimates. In row four, we use lagged employment as an additional regressor to control for the potential relevance of adjustment costs but find no noteworthy alterations to our estimates. Estimates from UBS wage data in the fifth row exhibit the same significance pattern as the UNIDO wage data, also with no significant effect in DEV and OIN. In contrast, OWW wage data in row six lead to a smaller sample and some alterations in the estimates for DEV and OIN wages. Relatively many missing OWW observations reduce the overlap with MIDI data below the overlap that UNIDO or UBS wage data can provide and make OWW results appear less reliable

Table 12: HOME-WAGE ELASTICITIES AT THE EXTENSIVE MARGIN

	Home wage change (1%), by regression specification							
	Stacking							Omnipr.
	UNIDO 98-01	UNIDO 00	UNIDO 98-01 AR(2)	UNIDO 98-01 lag y	UBS 98-01	OWW 98-01	UNIDO 98-01	UNIDO 98-01
Emplmt. chg. (%)	Ass. A	Ass. A	Ass. A	Ass. A	Ass. A	Ass. A	Ass. B	Ass. A
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
CEE	.794	.789	.791	.784	.720	.792	-12.840	.651
	(.043)***	(.156)***	(.111)***	(.045)***	(.094)***	(.144)***	(15.126)	(.127)***
DEV	.857	.783	.866	.827	1.048	.416	-25.420	.669
	(.098)***	(.159)***	(.276)***	(.101)***	(.369)***	(.329)	(27.013)	(.210)***
OIN	.629	.578	.750	.643	.582	1.140	-9.269	.501
	(.221)***	(.214)***	(.176)***	(.184)***	(.550)	(.690)*	(9.522)	(.157)***
WEU	.838	.843	.636	.820	1.206	.852	4.401	.612
	(.072)***	(.145)***	(.173)***	(.062)***	(.316)***	(.163)***	(4.064)	(.113)***
Obs.	1,654	326	1,654	1,654	1,628	1,467	1,654	96

Sources: MIDI and USTAN 1996 to 2001 (UNIDO, UBS and OWW wages).
Notes: Elasticities of home wage effects on foreign employment (first column of elasticity matrix) at the extensive margin. Standard errors from 200 bootstraps: * significance at ten, ** five, *** one percent. Locations: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

because region-wide wages are based on fewer observations. Nonparametric estimation under Assumption B does not yield statistically different estimates (row seven). The subsample of 96 omnipresent MNEs at the outcome stage is small and does not offer significant intensive-margin estimates (last row). Estimates for DEV and OIN are generally not statistically significant. This is consistent with the idea that there are no statistically relevant intensive-margin responses to wage shocks in remote locations.

Home-wage elasticities of foreign employment at the intensive margin are robust too, as Table 11 shows. Estimates on our benchmark sample (now in the first column) conform closely to several other specifications. In fact, our comments on the rows of Table 10 above apply equally to the columns of Table 11.

At the extensive margin, Table 12 documents that home-wage elasticities of foreign employment are (highly) significant in the parametric specifications (columns 1 through 6). Neither the restriction to the year-2000 cross section with a four-year selection lag, nor an AR(2) error specification for selection, nor the inclusion of lagged employment yield a significantly different elasticity estimate at any location. UBS wage data exhibit the same significance pattern as the UNIDO wage data, with highly significant effects at the extensive margin in CEE, DEV and WEU. In the relatively less complete OWW wage data, the elasticity point estimates in DEV and OIN become statistically insignificant. Nonparametric estimates of elasticities at the extensive margin are sample means of the first derivatives of our third-order polynomial series expansions. We compute the elasticities after dropping those outlier predictions for which the first-stage probability model would result in propensity scores outside the zero-one range. Nonparametric estimates for the extensive margin (column 7 of Table 12) are not statistically different from zero. Although the inclusion of nonparametric series terms in labor demand estimation yields more precise estimates of intensive margin coefficients (Tables 10 and 11 before), the series terms do not provide a precise estimate of the extensive margin itself. Point estimates for the extensive margin at omnipresent MNEs (last column) are statistically highly significant. Omnipresence is defined as presence in all world locations at the outcome stage. Taken together with the omnipresent MNEs' statistically insignificant responses at the intensive margin, this evidence is consistent with the view that omnipresent MNEs more heavily rely on the extensive margin, becoming omnipresent, because they do not significantly shift employment at the intensive margin.

In summary, robustness checks confirm the statistical plausibility of the benchmark estimates in Table 7 under parametric selectivity correction (Assumption A). In particular, estimates for DEV and OIN suggest that intensive-margin adjustment is weak but extensive-margin adjustment relatively strong at remote locations. Nonparametric estimates (Assumption B) are similar and highly significant at the intensive margin, but lack statistical significance at the extensive margin.

F Country groups by initial wage quartile

We turn to the robustness of our aggregate location definition by considering a different division of world regions: we split the world into the home country and four artificial locations defined by the quartiles of UNIDO manufacturing wages in the initial sample year 1996. We report estimated cross-wage elasticities at the two margins in Table G.3 in the Appendix. Four striking facts emerge. First, on-diagonal entries remain significantly negative and magnitudes off the diagonal exhibit substitutability when statistically significant. Second, quartiles 1 and 3, which happen to contain more distant countries from Germany, do not show statistically significant foreign-wage elasticities on home employment at the intensive margin, similar to the distant DEV and OIN locations before. These two facts corroborate our findings for the aggregate locations. Third, estimates at the selection margin show no variability off the diagonal for any given column (reflecting merely the observed wage bill share in eq. (4)). An economic interpretation is that the selection margin is not well defined for the artificial four-quartile regions that lack geographical and institutional coherence. Fourth, more off-diagonal entries of intensive-margin estimates are statistically significant than under our aggregate location definition. An economic interpretation is that intensive-margin substitutability cuts across the artificial four-quartile regions more frequently than across the geographically and institutionally related aggregate locations. The latter two facts support our definitions of aggregate locations as more coherent.

G Evaluation of magnitudes

We turn to the economic importance of our estimates for multinational labor substitution. Our hypothetical experiment is a permanent change in the wage differential between home and foreign locations. How much larger would parent employment be if the wage gap to foreign locations narrowed? How much smaller would affiliate employment be? We use the home-wage elasticities of foreign employment and the foreign-wage elasticities of home employment from Table 7. These estimates reflect the mean MNE's labor-demand response (the mean MNE in the stacked sample has propensities of presence abroad as in the first row of Table 4). We multiply the elasticity estimates with the workforce totals in Table 1 and obtain the implied employment changes from one-percent increases in wages by margin.

Table 13 displays the results. As the middle row shows, only the extensive margin is statistically relevant for labor substitution in all four foreign regions. Imprecise intensive-margin estimates for the remote regions DEV and OIN (first row) coincide with statistically insignificant total elasticity estimates (last row) for DEV and OIN. We compare responses in the close and coherent regions CEE and WEU. The bulk of the employment response in CEE occurs at the extensive margin, while the extensive margin accounts for around two-thirds of employment shifts in WEU. A one percent smaller wage gap between Germany and CEE is associated with around 380 more jobs at German parents and 2,000 less jobs at affiliates in CEE overall. CEE affiliates tend to have smaller work forces and, arguably, lower labor productivity than German establishments so that CEE employment is more sensitive to home wage changes than home employment to foreign wages. At

Table 13: COUNTERFACTUAL EMPLOYMENT EFFECTS OF A ONE-PERCENT REDUCTION IN THE HOME-FOREIGN WAGE GAP

	Permanent wage gap reduction by one percent between Home and			
Employment effect	CEE	DEV	OIN	WEU
on margin	(1)	(2)	(3)	(4)
Home[a] *total*	374	-40	1,214	2,820
	(75)***	(116)	(1077)	(901)***
Foreign[b] *extensive*	-1,951	-2,850	-2,008	-3,306
	(107)***	(326)***	(706)***	(284)***
Foreign[b] *total*	-2,046	271	-3,673	-4,979
	(394)***	(1560)	(3794)	(1574)***

Sources: Own calculations based on selectivity corrected translog estimates for 1,654 German manufacturing MNEs and their majority-owned foreign manufacturing affiliates in MIDI and USTAN between 1996 and 2001 (UNIDO wages).
Notes: Point estimates from parametric selectivity correction (Assumption A, Table 7) multiplied by employment in 2000 (Table 1). Standard errors from 200 bootstraps: * significance at ten, ** five, *** one percent. Home (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

[a]Gap reducing foreign wage increases (by one percent).

[b]Gap reducing home wage reduction (by one percent).

the margin, the labor substitution effect of 380 more home jobs for a one-percent higher CEE wage is considerably smaller than the effect for WEU, where a foreign wage increase by a percent brings 2,820 counterfactual jobs back to Germany. In absolute magnitude, however, a closing of the CEE-HOM wage gap by half *at constant elasticities* results in larger employment effects than a reduction of the WEU-HOM wage gap by half. The population-weighted mean UNIDO wages in CEE are 9.9 percent of the German level in 2000 (population-weighted mean OWW wages in CEE are 9.8 percent). If the estimated elasticities of substitution remained constant at all wage levels, an increase in CEE wages by 450% ($= [(1-.099)/2]/.099$) to reduce the wage gap vis-à-vis Germany by half in 2000 would bring 170,000 ($= 380 \cdot 450$) counterfactual manufacturing jobs to Germany—around an eighth of the estimated home employment at German manufacturing MNEs in 2000 (Table 1).[28] The UNIDO wage level in WEU is 78.6 percent of that in Germany so that an increase in WEU wages by 14% to cut the gap by half would attract only 40,000 counterfactual manufacturing jobs to the German MNE plants.[29]

[28]If international wage gaps shrink at a similar rate as per capita GDP converges to steady state and Germany is close to its steady state, the CEE-German wage gap would take around 35 years to contract to half its present size (Barro and Sala-i-Martin 1992).

[29]Recent survey evidence for 471 German manufacturing plants in 2001, 2003 and 2006 is consistent with employ-

IV Conclusion

The idea that multinational enterprises (MNEs) substitute jobs at home for foreign employment is widely espoused in public discourse. But economic research on MNE labor demand across locations has found weak or no evidence of job substitution. We unify two separate branches of the literature—one on predictions of MNEs' location choices, and one on labor substitutability across established MNE locations—into an integrated econometric model that embeds location selectivity into labor-demand estimation. In our framework, multinational labor demand responds to wage differentials across locations both at the extensive margin, when an MNE expands into foreign locations, and at the intensive margin, when an MNE reallocates jobs across existing affiliates. Our novel estimation strategy detects a complementarity bias in some elasticities from incomplete estimation models, but a substitutability bias in other cases.

Empirical evidence on German manufacturing MNEs shows that firms change multinational presence only infrequently and hardly alter their number of affiliates within regions. These scant changes to multinational presence at the extensive margin are associated with salient labor demand effects in response to permanent wage differentials. The extensive margin plays a crucial role for the resulting employment shifts. Equality between the intensive and the total elasticity of labor substitution is soundly rejected for every foreign location. For distant overseas locations, statistically significant employment responses occur only at the extensive margin.

These employment responses reflect MNEs' global employment decisions, given their product-market shares. In industry equilibrium, at least two additional employment effects might arise. First, MNEs with cost or market-access advantages after a foreign expansion might gain product-market shares and consequently employment. But, second, competitors at whose expense MNEs gain product-market shares might lose employment. Measuring the net employment effect in market equilibrium remains a task for future research. According estimation of product-market responses will require representative data on MNEs' national competitors and involve more restrictive assumptions than our generic model of a cost minimizing firm needs. Naturally, however, employment reallocation within MNEs, at the intensive and the extensive margins, will remain an integral aspect of industry equilibrium.

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ment shifts back to Germany: within four to five years after a foreign expansion, between one in six and one in four MNEs fully reverse the expansion (Fraunhofer ISI, PI Bulletin no. 45/2008).

References

Barba Navaretti, Giorgio and Davide Castellani, "Do Italian Firms Improve Their Performance at Home by Investing Abroad?," in Steven Brakman and Harry Garretsen, eds., *Foreign Direct Investment and the Multinational Enterprise*, CESifo Seminar Series, Cambridge and London: MIT Press, 2008, chapter 9, pp. 199–224.

Barro, Robert J. and Jong Wha Lee, "International Data on Educational Attainment: Updates and Implications," *Oxford Economic Papers*, July 2001, *53* (3), 541–63.

_ **and Xavier Sala-i-Martin**, "Convergence," *Journal of Political Economy*, April 1992, *100* (2), 223–51.

Barten, A. P., "Maximum Likelihood Estimation of a Complete System of Demand Equations," *European Economic Review*, Autumn 1969, *1* (1), 7–73.

Becker, Sascha O. and Marc-Andreas Muendler, "The Effect of FDI on Job Security," *The B.E. Journal of Economic Analysis & Policy: Advances*, April 2008, *8* (1), Article 8, pp. 1–45.

_ **, Karolina Ekholm, Robert Jäckle, and Marc-Andreas Muendler**, "Location Choice and Employment Decisions: A Comparison of German and Swedish Multinationals," *Review of World Economics/Weltwirtschaftliches Archiv*, December 2005, *141* (4), 693–731.

Bernard, Andrew B., J. Bradford Jensen, and Peter K. Schott, "Importers, Exporters, and Multinationals: A Portrait of Firms in the U.S. that Trade Goods," in Timothy Dunne, J. Bradford Jensen, and Mark J. Roberts, eds., *Producer Dynamics: New Evidence from Micro Data*, Vol. 68 of *Studies in Income and Wealth*, Chicago: University of Chicago Press, May 2009, chapter 14, pp. 513–552.

Blonigen, Bruce A., Ronald B. Davies, and Keith Head, "Estimating the Knowledge-Capital Model of the Multinational Enterprise: Comment," *American Economic Review*, June 2003, *93* (3), 980–94.

Braconier, Henrik and Karolina Ekholm, "Swedish Multinationals and Competition from High- and Low-Wage Locations," *Review of International Economics*, August 2000, *8* (3), 448–61.

Buch, Claudia M., Jorn Kleinert, Alexander Lipponer, and Farid Toubal, "Determinants and Effects of Foreign Direct Investment: Evidence from German Firm-Level Data," *Economic Policy*, January 2005, *41*, 51–98.

Carr, David L., James R. Markusen, and Keith E. Maskus, "Estimating the Knowledge-Capital Model of the Multinational Enterprise," *American Economic Review*, June 2001, *91* (3), 693–708.

Chen, Songnian and Shakeeb Khan, "Semiparametric Estimation of a Heteroskedastic Sample Selection Model," *Econometric Theory*, December 2003, *19* (6), 1040–64.

Das, Mitali, Whitney K. Newey, and Francis Vella, "Nonparametric Estimation of Sample Selection Models," *Review of Economic Studies*, January 2003, *70* (1), 33–58.

de Fontenay, Catherine C. and Joshua S. Gans, "Organizational Design and Technology Choice under Intrafirm Bargaining: Comment," *American Economic Review*, March 2003, *93* (1), 448–55.

Debaere, Peter, Hongshik Lee, and Joonhyung Lee, "Does Where You Go Matter? The Impact of Outward Foreign Direct Investment on Multinationals' Employment at Home," *CEPR Discussion Paper*, July 2006, *5737.*

Desai, Mihir A., C. Fritz Foley, and James R. Hines, "Domestic Effects of the Foreign Activities of U.S. Multinationals," *American Economic Journal: Economic Policy*, February 2009, *1* (1), 181–203.

Deutsche Bundesbank, "The Methodological Basis of the Deutsche Bundesbank's Corporate Balance Sheet Statistics," *Monthly Report*, October 1998, *1998* (10), 49–64.

Devereux, Michael P. and Rachel Griffith, "Taxes and the Location of Production: Evidence from a Panel of U.S. Multinationals," *Journal of Public Economics*, June 1998, *68* (3), 335–67.

Dhrymes, Phoebus J., "Equivalence of Iterative Aitken and Maximum Likelihood Estimators for a System of Regression Equations," *Australian Economic Papers*, June 1971, *10* (16), 20–24.

Disdier, Anne Celia and Thierry Mayer, "How Different Is Eastern Europe? Structure and Determinants of Location Choices by French Firms in Eastern and Western Europe," *Journal of Comparative Economics*, June 2004, *32* (2), 280–96.

Eakin, B. Kelly, Daniel P. McMillen, and Mark J. Buono, "Constructing Confidence Intervals Using the Bootstrap: An Application to a Multi-Product Cost Function," *Review of Economics and Statistics*, May 1990, *72* (2), 339–44.

Egger, Peter and Michael Pfaffermayr, "The Counterfactual to Investing Abroad: An Endogenous Treatment Approach of Foreign Affiliate Activity," *University of Innsbruck Working Paper*, April 2003, *02.*

Ekholm, Karolina and Marie Hesselman, "The Foreign Operations of Swedish Manufacturing Firms: Evidence from a Survey of Swedish Multinationals 1998," *Research Institute of Industrial Economics (IUI) Working Paper*, 2000, *540.*

Feenstra, Robert C. and Gordon H. Hanson, "The Impact of Outsourcing and High-Technology Capital on Wages: Estimates for the United States, 1979-1990," *Quarterly Journal of Economics*, August 1999, *114* (3), 907–40.

Freeman, Richard B. and Remco H. Oostendorp, "The Occupational Wages around the World Data File," *International Labour Review*, 2001, *140* (4), 379–401.

Hanson, Gordon H., Raymond J. Mataloni, and Matthew J. Slaughter, "Vertical Production Networks in Multinational Firms," *Review of Economics and Statistics*, November 2005, *87* (4), 664–78.

Harrison, Ann E. and Margaret S. McMillan, "Outsourcing Jobs? Multinationals and U.S. Employment," *NBER Working Paper*, July 2006, *12372*.

Head, Keith and Thierry Mayer, "Market Potential and the Location of Japanese Investment in the European Union," *Review of Economics and Statistics*, November 2004, *86* (4), 959–72.

Heckman, James J., "Sample Selection Bias as a Specification Error," *Econometrica*, January 1979, *47* (1), 153–61.

Jäckle, Robert and Georg Wamser, "Going Multinational: What Are the Effects on Home Market Performance?," *German Economic Review*, forthcoming.

Klein, Roger W. and Richard H. Spady, "An Efficient Semiparametric Estimator for Binary Response Models," *Econometrica*, March 1993, *61* (2), 387–421.

Koebel, Bertrand M., "Exports and Labor Demand: Searching for Functional Structure in Multi-Output Multi-Skill Technologies," *Journal of Business and Economic Statistics*, January 2006, *24* (1), 91–103.

Konings, Jozef and Alan Patrick Murphy, "Do Multinational Enterprises Relocate Employment to Low-Wage Regions? Evidence from European Multinationals," *Review of World Economics/Weltwirtschaftliches Archiv*, July 2006, *142* (2), 267–86.

Kotz, Samuel, N. Balakrishnan, and Norman L. Johnson, *Continuous Multivariate Distributions*, 2nd ed., Vol. 1 of *Wiley Series in Probability and Statistics*, New York: Wiley, 2000.

Lee, Lung Fei, "Semiparametric Maximum Likelihood Estimation of Polychotomous and Sequential Choice Models," *Journal of Econometrics*, February 1995, *65* (2), 381–428.

_ **and Mark M. Pitt**, "Microeconometric Demand Systems with Binding Nonnegativity Constraints: The Dual Approach," *Econometrica*, September 1986, *54* (5), 1237–42.

Lipponer, Alexander, "A "New" Micro Database for German FDI," in Heinz Herrmann and Robert Lipsey, eds., *Foreign Direct Investment in the Real and Financial Sector of Industrial Countries*, Berlin: Springer, 2003, pp. 215–44.

Marin, Dalia, "A Nation of Poets and Thinkers—Less so with Eastern Enlargement? Austria and Germany," *CEPR Discussion Paper*, 2004, *4358*.

Markusen, James R., "Trade in Producer Services and in Other Specialized Intermediate Inputs," *American Economic Review*, March 1989, *79* (1), 85–95.

Neary, J. Peter and Kevin W. S. Roberts, "The Theory of Household Behaviour under Rationing," *European Economic Review*, January 1980, *13* (1), 25–42.

Pan, Wei, "Akaike's Information Criterion in Generalized Estimating Equations," *Biometrics*, March 2001, *57* (1), 120–125.

Redding, Stephen and Anthony J. Venables, "Economic Geography and International Inequality," *Journal of International Economics*, January 2004, *62* (1), 53–82.

Riker, David A. and S. Lael Brainard, "U.S. Multinationals and Competition from Low Wage Countries," *NBER Working Paper*, 1997, *5959.*

Slaughter, Matthew J., "Production Transfer within Multinational Enterprises and American Wages," *Journal of International Economics*, April 2000, *50* (2), 449–72.

Stole, Lars A. and Jeffrey Zwiebel, "Intra-firm Bargaining under Non-binding Contracts," *Review of Economic Studies*, July 1996, *63* (3), 375–410.

_ **and** _ , "Organizational Design and Technology Choice under Intrafirm Bargaining," *American Economic Review*, March 1996, *86* (1), 195–222.

UBS, "Prices and Earnings: A Comparison of Purchasing Power Around the World," *ubs.com* Zürich 2003.

UNIDO, "INDSTAT3 2005 ISIC Rev. 2 User's Guide," *unido.org/doc/3531*, Vienna 2005.

Wolinsky, Asher, "A Theory of the Firm with Non-binding Employment Contracts," *Econometrica*, July 2000, *68* (4), 875–910.

Zellner, Arnold, "An Efficient Method for Estimating Seemingly Unrelated Regressions and Tests for Aggregation Bias," *Journal of the American Statistical Association*, June 1962, *57* (298), 348–68.

Appendix FOR ONLINE PUBLICATION

A Multiproduct translog cost function

Consider the short-run multiproduct translog function with quasi-fixed capital:[30]

$$\begin{aligned}
\ln C_{jt} \quad = \quad & \varphi + \sum_{n=1}^{L} \varphi_n^0 \ln q_{jt}^n + \sum_{\ell=1}^{L} \alpha_\ell \ln w_t^\ell + \sum_{n=1}^{L}\sum_{\ell=1}^{L} \mu_{\ell n} \ln q_{jt}^n \ln w_t^\ell \\
& + \frac{1}{2}\sum_{n=1}^{L}\sum_{\ell=1}^{L} \varphi_{\ell n}^1 \ln q_{jt}^n \ln q_{jt}^\ell + \frac{1}{2}\sum_{n=1}^{L}\sum_{\ell=1}^{L} \delta_{\ell n} \ln w_t^n \ln w_t^\ell \\
& + \sum_{n=1}^{L} \zeta_n^0 \ln k_{jt}^n + \sum_{n=1}^{L}\sum_{\ell=1}^{L} \zeta_{\ell n}^{11} \ln k_{jt}^n \ln q_{jt}^\ell \qquad\qquad \text{(A1)} \\
& + \sum_{n=1}^{L}\sum_{\ell=1}^{L} \kappa_{\ell n} \ln k_{jt}^n \ln w_t^\ell + \frac{1}{2}\sum_{n=1}^{L}\sum_{\ell=1}^{L} \zeta_{\ell n}^1 \ln k_{jt}^n \ln k_{jt}^\ell .
\end{aligned}$$

By Shepard's lemma, MNE j's demand for employment y_{jt}^ℓ is equal to $\partial C_{jt}/\partial w_t^\ell$ so that the wage bill share $s_{jt}^\ell \equiv w_t^\ell y_{jt}^\ell / C_{jt}$ at location ℓ becomes

$$s_{jt}^\ell = \frac{\partial C_{jt}/\partial w_t^\ell}{C_{jt}/w_t^\ell} = \alpha_\ell + \sum_{n=1}^{L} \left(\mu_{\ell n} \ln q_{jt}^n + \kappa_{\ell n} \ln k_{jt}^n + \delta_{\ell n} \ln w_t^n \right)$$

for $\ell = 1, \ldots, L$. We transform these L equations into L simultaneous labor demand functions by multiplying the dependent variable and all regressors with the observation-specific scalars C_{jt}/w_t^ℓ and obtain $y_{jt}^\ell = \partial C_{jt}/\partial w_t^\ell = s_{jt}^\ell C_{jt}/w_t^\ell$ as in eq. (1).

With L locations, there are $L(L-1)/2$ symmetry restrictions $\delta_{k\ell} = \delta_{\ell k}$ for any k, ℓ. Linear homogeneity in factor prices requires that $\sum_{\ell=1}^{L} \alpha_\ell = 1$ and that $\sum_{\ell=1}^{L} \mu_{\ell n} = \sum_{\ell=1}^{L} \kappa_{\ell n} = \sum_{\ell=1}^{L} \delta_{\ell n} = \sum_{\ell=1}^{L} \delta_{n \ell} = 0$ for all n. We impose these restrictions on intensive-margin estimation but do not constrain extensive-margin coefficients.

B Stacking

Eq. (1) requires treatment for locations of absence because outputs and capital inputs are missing where MNEs do not operate. Our maintained assumptions imply that stacking of observations is a

[30] Slaughter (2000) adds $\ln(k/q)$ terms to a version of (A1). Given the additive logarithmic structure, this is equivalent to an affine transformation of the parameter pairs (α_k, ζ_k) and $(\mu_{k,\ell}, \kappa_{k,\ell})$ because $\ln(k/q) = \ln k - \ln q$.

viable and attractive procedure.[31] Stacking means that we set regressors for locations of absence to zero. Stacking is easily implemented, improves efficiency, collapses the up to $2^{L-1}-1$ sets of estimates into one consistently estimated $(L-1)$-equation system, and provides a single $L \times L$ matrix of estimates for wage elasticities of regional labor demands.

More formally, stacking interacts the parameters in (1) with presence indicators: $\mu_{\ell n}=0$ when no output is produced at location n, and $\kappa_{\ell n}=\delta_{\ell n}=0$ when MNE j employs no factors at location n. Stacking is permissible under three natural assumptions in our framework: (i) all MNEs face the same sunk cost function $F^{\ell}_{j,t-\tau}$ conditional on prior presence (so that presence is mean independent of inputs); (ii) MNEs face an identical short-run cost function $C(\cdot)$ in all locations of presence (but not necessarily where absent) conditional on characteristics (so that a common parameter vector is justified); and (iii) the disturbances ϵ^{ℓ}_{jt} are uncorrelated across observations of MNEs i and j. To prevent any bias from stacking, we include a set of absence indicators $(\mathbf{1}-\mathbf{d}^{n\neq\ell}_{jt})$ in the outcome equation. Absence indicators control for shadow inputs. To check robustness of the stacking procedure, we repeat estimation for the subsample of omnipresent MNEs that operate affiliates in all locations.

C Parametric selection correction

Given our parametric cost function, a parametric approach to selectivity is a natural benchmark. Plausible distributional assumptions permit individual Heckman (1979) corrections location by location.[32] Consider linear selection predictions $H(\mathbf{z}_{j,t-\tau}) = \mathbf{z}_{j,t-\tau}\gamma^{\ell}$ and jointly normally distributed disturbances $(\epsilon^{k}_{jt}, \eta^{\ell}_{j,t-\tau})$ so that a probit model describes the choice of presence (6).

The correlation between ϵ^{n}_{jt} and $\eta^{\ell}_{j,t-\tau}$ across separate locations $n \neq \ell$ is crucial for estimation of outcomes (2). Our data reject independence of ϵ^{n}_{jt} and $\eta^{\ell}_{j,t-\tau}$.[33] To specify the correlation

[31] Estimation of separate equation systems for all possible presence patterns is plagued by dimensionality: potential presence in up to $L-1$ locations outside home means that there are up to $2^{L-1}-1$ regional presence patterns. Lee and Pitt (1986) propose an estimator related to Neary and Roberts's (1980) shadow price approach. Koebel (2006) conducts Box-Cox transformations on inputs.

[32] For multivariate selectivity, an extension of the univariate Heckman (1979) estimator has a complicated form (conditional moments of multivariate normal distributions have no known closed form for multiple truncations, see Kotz, Balakrishnan and Johnson (2000)). Simulated maximum-likelihood would be a viable technique but requires joint multivariate normality, which we prefer to relax in nonparametric estimation.

[33] SUR estimation of the outcome equations shows that ϵ^{n}_{jt} and ϵ^{ℓ}_{jt} correlate so that ϵ^{n}_{jt} and $\eta^{\ell}_{j,t-\tau}$ must be correlated because ϵ^{ℓ}_{jt} and $\eta^{\ell}_{j,t-\tau}$ are correlated.

structure, we depart from the idea that selection disturbances include both location-specific parts such as, for example, surprising changes to profit repatriation policies in the host country and include MNE-specific parts such as idiosyncratic shocks to a firm's sunk entry costs. Changes to host-country repatriation policies affect the entry decision. But once the MNE operates in the host country, it minimizes costs irrespective of entry-related host-country shocks. So, we consider it plausible to assume that there is an MNE-specific, location-independent component e_{jt} to the selection shock $\eta^n_{j,t-\tau}$ and that the labor-demand shock ϵ^ℓ_{jt} correlates with the selection shock $\eta^n_{j,t-\tau}$ elsewhere only through the MNE-specific component e_{jt}. The assumption is not rejected in our data. Note that, under this assumption, cost function disturbances do covary with entry shocks across locations, but only through an MNE-specific component.

Assumption A *The disturbances ϵ^n_{jt} and $\eta^\ell_{j,t-\tau}$ are multivariate normally distributed with $\epsilon^n_{jt} = \lambda\, e_{jt} + \pi^n_\epsilon\, v^n_{jt}$ and $\eta^\ell_{j,t-\tau} = \sqrt{1-\nu}\, e_{jt} + \sqrt{\nu}\, u^\ell_{jt}$, where $\nu \in [0,1]$ and the standard normal variables $e_{jt}, u^\ell_{jt}, v^n_{jt}$ are independent of $\mathbf{x}^m_{jt}$ and $\mathbf{z}_{j,t-\tau}$ for all ℓ, m, n.*

Any normally distributed random variable can be decomposed into an affine function of standard normal variables. Assumption A does this. Under Assumption A, the variances and covariances of the selection shocks are $\sigma^{\ell\ell}_\eta = 1$, as is common for probit, and $\sigma^{n\ell}_\eta = 1-\nu$. The variances and covariances of the labor demand shocks are $\sigma^{\ell\ell}_\epsilon = \lambda^2 + (\pi^{\ell\ell}_\epsilon)^2$ and $\sigma^{n\ell}_\epsilon = \lambda^2$. And the covariances between the selection shock in location n and the demand shock in location ℓ are $\sigma^{n\ell}_{\eta\epsilon} = \lambda$. So, cost function disturbances do correlate with entry-relevant policy shocks across locations, but only through an MNE-specific shock. The assumption accommodates potential serial correlation in location selection, defining $u^\ell_{jt} \equiv \sum_\varsigma \alpha^u_\varsigma \tilde{u}^\ell_{j,t-\varsigma}$. Assumption A is testable. We obtain estimates of $\sigma^{n\ell}_\eta = 1-\nu$ from multivariate probit estimation (on the same set of regressors as in Table 5) and use a χ^2-test for their equality. We fail to reject equality.

Intuitively, all selection-related information that is relevant for labor demand at any location ℓ is fully contained in the single presence indicator d^ℓ_{jt}, which is as informative about $\eta^\ell_{j,t-\tau}$ as any other location indicator. So, location-by-location correction for selectivity is permissible.

Lemma 1 *Independent selection correction for L locations identifies $\mathbf{x}^\ell_{jt}\beta^\ell$ and $m^\ell\big(\Pr^\ell(\mathbf{z}_{j,t-\tau})\big)$ if Assumption A holds.*

Proof. Denote the standard normal density and distribution functions with $\phi(\cdot)$ and $\Phi(\cdot)$. Under Assumption A, the marginal likelihood function is

$$g(y_{jt}^{\ell}|\mathbf{x}_{jt}^{\ell},\mathbf{z}_{j,t-\tau}) = \frac{\phi\left((y_{jt}^{\ell}-\mathbf{x}_{jt}^{\ell}\beta^{\ell})/\sigma_{\epsilon}^{\ell}\right)}{\sigma_{\epsilon}^{\ell}\,\Phi(\mathbf{z}_{j,t-\tau}\gamma^{\ell})}\cdot\Phi\!\left(\frac{\rho_{\eta\epsilon}^{\ell\ell}(y_{jt}^{\ell}-\mathbf{x}_{jt}^{\ell}\beta^{\ell})+\mathbf{z}_{j,t-\tau}\gamma^{\ell}}{\sigma_{\epsilon}^{\ell}\,(1-\rho_{\eta\epsilon}^{\ell\ell})^{1/2}}\right),$$

after concentrating out u_{jt}^{ℓ} and v_{jt}^{ℓ}, where $\sigma_{\epsilon}^{\ell}=\sqrt{\sigma_{\epsilon}^{\ell\ell}}=\sqrt{\lambda^2+(\pi_{\epsilon}^{\ell\ell})^2}$ and $\rho_{\eta\epsilon}^{\ell\ell}=\sigma_{\eta\epsilon}^{\ell\ell}/\sigma_{\epsilon}^{\ell}=\lambda/\sqrt{\lambda^2+(\pi_{\epsilon}^{\ell\ell})^2}$. This is the likelihood function for independent Heckman (1979) correction location by location, where $m^{\ell}\big(\Pr^{\ell}(\mathbf{z}_{j,t-\tau})\big)=\beta_{\Lambda}^{\ell}\Lambda_{jt}^{\ell}(\mathbf{z}_{j,t-\tau}\gamma^{\ell})$ and $\beta_{\Lambda}^{\ell}=\rho_{\eta\epsilon}^{\ell\ell}\sigma_{\epsilon}^{\ell}$ is the coefficient on the selectivity hazard $\Lambda_{jt}^{\ell}(\mathbf{z}_{j,t-\tau}\gamma^{\ell})$ (the inverse of the Mills ratio) in the outcome equation. ■

Under Heckman (1979) correction (Assumption A), the extensive-margin term in (5) simplifies to $\beta_{\Lambda}^{\ell}\,\Delta_{jt}^{\ell}\cdot\gamma_{w^n}^{\ell}\cdot w_t^{\ell}w_t^{n}/C$, where $\gamma_{w^n}^{\ell}$ is the wage coefficient in the selection equation, β_{Λ}^{ℓ} is the coefficient on the selectivity hazard in the outcome equation, and Δ_{jt}^{ℓ} is the first derivative of the selectivity hazard Λ_{jt}^{ℓ} (the inverse of the Mills ratio) with respect to its argument, $\Delta_j^{\ell}(\mathbf{z}_{j,t-\tau}\gamma^{\ell})\equiv\Lambda_j^{\ell}(\mathbf{z}_{j,t-\tau}\gamma^{\ell})[\Lambda_j^{\ell}(\mathbf{z}_{j,t-\tau}\gamma^{\ell})-\mathbf{z}_{j,t-\tau}\gamma^{\ell}]$. Because $\Delta_j^{\ell}(\cdot)\in(0,1)$, the sign of the log wage effect on the wage bill at the extensive margin is the sign of the product $\gamma_{w^n}^{\ell}\beta_{\Lambda}^{\ell}$ (the coefficients on the two stages of estimation).

D Nonparametric selection correction

To establish identification, consider the following deviations from the truth: $\Delta\xi^{\ell}(\mathbf{x}_{jt}^{\ell})\equiv\mathbf{x}_{jt}^{\ell}(\hat{\beta}^{\ell}-\beta^{\ell})$ and $\Delta m^{\ell}(\mathbf{P}_{jt})\equiv\hat{m}^{\ell}(\mathbf{P}_{jt})-m^{\ell}(\mathbf{P}_{jt})$, where hats denote estimates of the true (not hatted) functions.

Assumption B formally states one set of sufficient conditions for identification.

Assumption B

(i) $\mathbb{E}[\epsilon_{jt}^{\ell}\,|\,d_{jt}^{\ell}=1,\mathbf{z}_{j,t-\tau}]=m^{\ell}(\mathbf{P}_{jt})$ *and* $\mathbb{C}ov(\epsilon_{jt}^{\ell},\eta_{j,t-\tau}^{k})=0$ *for* $k\neq\ell$*,*

(ii) $\Pr(\Delta\xi^{\ell}(\mathbf{x}_{jt}^{\ell})+\Delta m^{\ell}(\mathbf{P}_{jt})=0|d_{jt}^{\ell}=1)=1$ *implies that* $\Delta\xi^{\ell}(\mathbf{x}_{jt}^{\ell})$ *is constant,*

(iii) $\nabla_{\mathbf{z}_{j,t-\tau}}\mathbf{P}_{jt}\neq\mathbf{0}$ *with probability one,*

for $\ell=1,\ldots,L$*.*

Part (i) posits that the conditional expectation of the labor demand disturbance at location ℓ is a function of the propensity scores of presence at any location $k = 1, \ldots, L$. So, in the regression of observed labor demand y^ℓ_{jt} on $\mathbf{x}^\ell_{jt}\beta^\ell$ and $m^\ell(\mathbf{P}_{jt})$, $\mathbf{x}^\ell_{jt}\beta^\ell$ is a separate additive component. This specification applies nonparametric selectivity correction with a single outcome equation (but multiple selection thresholds) in Das et al. (2003) to the multivariate outcome case.[34] The generalization to simultaneous location selection (multivariate selectivity) comes at a price. To maintain identifying restrictions similar to Das et al. (2003), we need to assume cross-equation independence in the selection disturbance conditional on observable variables.

Part (ii) is the same identification condition as in Das et al. (2003) and implies that $\mathbf{P}_{jt}$ (which enters $m^\ell(\mathbf{P}_{jt})$) depends on variables in $\mathbf{z}_{j,t-\tau}$ that are not in $\mathbf{x}^\ell_{jt}\beta^\ell$. Otherwise, a regression of y^ℓ_{jt} on $\mathbf{x}^\ell_{jt}\beta^\ell$ leaves $\Delta\xi^\ell(\mathbf{x}^\ell_{jt}) = m^\ell(\mathbf{P}_{jt})$ and $\Delta m^\ell(\mathbf{P}_{jt}) = -m^\ell(\mathbf{P}_{jt})$ indeterminate—a violation of (ii). In our context, parent-firm characteristics and competitor-level host-country characteristics are among the $\mathbf{z}_{j,t-\tau}$ predictors of presence but not related to the labor-specific part of the cost function other than through wages themselves. The rank condition (iii) requires that the information set $\mathbf{z}_{j,t-\tau}$ predicts the propensity score.

Lemma 2 *If Assumption B holds and if $m^\ell(\mathbf{P}_{jt})$ and $\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau})$ are continuously differentiable and have continuous distribution functions almost everywhere, then $\mathbf{x}^\ell_{jt}\beta^\ell$ and $m^\ell(\mathbf{P}_{jt})$ are identified up to additive constants.*

Proof. In any observationally equivalent model it must be the case that the observed outcome satisfies $\mathbb{E}[y^\ell_{jt} \,|\, \mathbf{x}^\ell_{jt}, \mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}] = \mathbf{x}^\ell_{jt}\hat{\beta}^\ell + \hat{m}^\ell(\mathbf{P}_{jt})$ for some $\mathbf{x}^\ell_{jt}\hat{\beta}^\ell$ and $\hat{m}^\ell(\mathbf{P}_{jt})$. Equivalently, deviations from the truth $\Delta\xi^\ell(\mathbf{x}^\ell_{jt}) + \Delta m^\ell(\mathbf{P}_{jt}) = 0$. This identity must be differentiable with respect to $\mathbf{x}^\ell_{jt}$ and $\mathbf{z}_{j,t-\tau}$ by continuous differentiability of $m^\ell(\mathbf{P}_{jt})$ and $\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau})$. So,

$$\begin{aligned}\nabla_{\mathbf{x}^\ell_{jt}}\Delta\xi^\ell(\mathbf{x}^\ell_{jt}) &= \mathbf{0},\\ (\nabla_{\mathbf{P}_{jt}}\Delta m^\ell(\mathbf{P}_{jt}))\cdot\nabla_{\mathbf{z}_{j,t-\tau}}\mathbf{P}_{jt} &= \mathbf{0}.\end{aligned}$$

The first equation implies that $\Delta\xi^\ell(\mathbf{x}^\ell_{jt}) = \mathbf{x}^\ell_{jt}(\hat{\beta}^\ell - \beta^\ell) = c_1$ for a constant c_1 and $\mathbf{x}^\ell_{jt}\beta^\ell$ is

[34] A semiparametric alternative would be the Lee (1995) estimator, a multivariate extension to Klein and Spady's (1993) semiparametric maximum-likelihood estimator. Lee (1995) partitions the covariates $\mathbf{z}_{j,t-\tau}$ to appear in $H(\mathbf{z}_{j,t-\tau})$ through multiple indexes. Note, however, that in our context the information set $\mathbf{z}_{j,t-\tau}$ includes location selection predictors from every world region; so there is no natural subpartition. A nonparametric estimator for $H(\mathbf{z}_{j,t-\tau})$ accommodates the multiple-index case and simultaneous selection into more than one location.

identified up to this constant. By $\nabla_{\mathbf{z}_{j,t-\tau}}\mathbf{P}_{jt} \neq \mathbf{0}$, the second equation implies that $\Delta m^\ell(\mathbf{P}_{jt}) = \hat{m}^\ell(\mathbf{P}_{jt}) - m^\ell(\mathbf{P}_{jt}) = c_2$ for a constant c_2 and $m^\ell(\mathbf{P}_{jt})$ is identified up to that constant. ■

Under nonparametric location selection (Assumption B) and polynomial series estimation, the derivatives of $m^\ell(\cdot)$ and $\Pr^\ell_{jt}$ at the extensive margin are the marginal effects on the polynomial terms $\nabla_{\mathbf{P}_{jt}} m^\ell(\mathbf{P}_{jt}) \cdot \nabla_{w^n_{t-\tau}}\mathbf{P}_{jt} \cdot w^\ell_t w^n_t / C$, which we evaluate at the sample mean.

E Modelling Wage Endogeneity

Cross-wage elasticities (4) were derived in the context of competitive labor markets. MNEs, however, pay wage premia over local competitors. Suggested reasons include relatively skilled workforces and rent sharing through efficiency wages or bargaining.

Under wage bargaining, cross-wage elasticities (4) remain consistent with departures from competitive labor markets. Stole and Zwiebel (1996a, 1996b), for instance, consider bargaining between a firm and its individual workers, whose contracts cannot bind them to the firm. Their model relates bargaining outcomes to a firm's individual profitability and can explain within-industry wage differences between firms, such as mark-ups at MNEs relative to local competitors, if there are fixed hiring costs at wage-bargaining firms. A wage-bargaining firm's cost function does not necessarily exhibit first-degree homogeneity in paid wages. But, in line with our translog cost specification where we use location-wide median wages as outside wages, the wage-bargaining firm's cost function is homogeneous of degree one in location-specific reservation wages.

To establish homogeneity of degree in location wages, note that the first-order condition in Stole and Zwiebel (1996a, 1996b) for single-product firms requires that, at the optimal employment level $\tilde{n}$, realized profits are equal to average profits over all putative inframarginal workforce sizes

$$\pi(n,k) = p\,q(n,k) - \underline{w}n - \underline{r}k = (1/\tilde{n}) \int_0^{\tilde{n}} \pi(s,k)ds \equiv \tilde{\pi}(\tilde{n},k),$$

where $\underline{w}$ and $\underline{r}$ are reservation factor prices. Since optimal profits $\pi(\tilde{n},k)$ are homogeneous of degree one in reservation prices by this first-order condition (an instance of the envelope theorem), the cost function is homogeneous of degree one in reservation wages. Similarly, Shepard's lemma holds for the reservation wage.

The consequences of wage bargaining for labor demand are theoretically ambiguous when contracts are non-binding (Stole and Zwiebel 1996a, Wolinsky 2000, de Fontenay and Gans 2003). To capture potential employment distortions, whatever their direction, we include MNE-specific wage residuals beyond the main location-specific median wages in our labor demand system.

We use the predicted log wage residual ψ^{ℓ}_{jt} from a reduced form regression, mirroring the selection equation, to control for potential bias that could arise from omitting an MNE-specific wage premium or discount relative to the industry-wide median wage at the location:

$$\ln w^{\ell}_{jt} = W(\mathbf{z}_{j,t-\tau}) + \psi^{\ell}_{jt},$$

where w^{ℓ}_{jt} is the MNE's paid wage at location j, $W(\cdot)$ mirrors the functional form of the location-selection equation (linear in the Heckman model), and $\mathbf{z}_{j,t-\tau}$ is the vector of selection predictors. We include the set of predicted residual wages ψ^{ℓ}_{jt} for all locations with an MNE's presence as additional regressors in outcome eq. (2) on the second stage. By construction, the log MNE-wage residuals are orthogonal to the propensity score. So any wage variation associated with the propensity score of presence is assigned to the extensive margin, as our selection model requires.

Including the estimated log wage residuals in labor-demand equations addresses concerns with profit-related pay and unobserved workforce heterogeneity. As suggested by firm-level wage bargaining described above, more productive MNEs may share rents with their workforces across locations and the MNEs' profitability may covary with industry-median wages at those locations. Alternatively, more productive MNEs may employ more skilled workforces, which can be associated with industry-median wages if the productivity dispersion is industry dependent. So, unobserved MNE productivity could bias our employment estimation unless firm-specific wage residuals are included alongside industry-wide wages. The firm-specific wage residuals serve as controls only. In line with the structural MNE model, estimation of cross-wage elasticities from labor-demand outcomes exclusively rests on the industry-wide median wage coefficients ($\delta_{\ell n}$).

F Data

F.1 Currency conversion and deflation

We deflate parent variables with the German consumer price index and deflate affiliate variables with country-level consumer price indices (from the IMF's International Financial Statistics).[35] CPI series are available for a broader set of countries than producer or wholesale price series. CPIs properly reflect the opportunity costs for investors who are the beneficiaries of firms' profit maximization. We re-base CPI deflation factors to unity at year end 1998 and transform foreign currency values to their EUR equivalents in December 1998 in order to remove nominal exchange rate fluctuations. December 1998 is the mid point in time for our 1996-2001 sample. Introduction of the euro in early 1999 makes December 1998 a natural reference date.

In BuBa's original MIDI data, all information on foreign affiliates is reported in German currency using the exchange rate at the closing date of the foreign affiliate's balance sheet. Concretely, we apply the following conversion to all financial variables, including the physical capital stock (fixed assets). Deutschmark (DEM) figures are transformed into EUR at the rate 1/1.95583 (the conversion rate at euro inception in 1999). (i) We use the market exchange rate on the end-of-month day closest to an affiliate's balance sheet closing date to convert the DEM or EUR figures into local currency for every affiliate. This reverses the conversion applied to the questionnaires at the date of reporting. (ii) A CPI factor for every country deflates the foreign-currency financial figures to the December-1998 real value in local currency. (iii) For each country, the average of all end-of-month exchange rates vis-à-vis the DEM or EUR between January 1996 and December 2001 is used as a proxy for purchasing power parity of foreign consumption baskets relative to the DEM or EUR. All deflated local-currency figures are converted back to DEM or EUR using this purchasing-power proxy.

F.2 Wages

Our main estimation sample uses sectoral manufacturing wages by country between 1996 and 2001 from the UNIDO Industrial Statistics Database at the 3-digit ISIC level, Rev. 2 (UNIDO 2005). The

[35] We use the CPI in the currency-issuing country whenever a country's CPI is not available from IFS but the main currency is issued elsewhere. We use current exchange rates and the German price deflator whenever foreign price deflators are missing or period-average exchange rate information is incomplete.

UNIDO measure of annual sectoral wage bills includes all payments to workers at establishments in the reference sector and year (wages and salaries, remuneration for time not worked, bonuses and gratuities, allowances, and payments in kind; but excludes contributions to social security, pensions, insurance, severance and termination pay). We divide the sectoral wage bill by the sectoral number of workers and employees. The UNIDO data cover 109 countries and result in the largest overlap with MIDI observations.

For robustness checks, we use wage data collected by the Swiss commercial bank UBS for metropolitan areas around the world in 1994, 1997, 2000 and 2003 (UBS 2003). We linearly interpolate UBS wages between survey years to cover our sample period 1996-2001. UBS surveyed approximately 70 cities during the second quarter of 1994, 1997 and 2000, and during the first quarter of 2003. Questionnaires request detailed information on wage components, wage deductions and working hours across thirteen occupations. UBS converts wage figures into U.S. dollars and smoothes the effect of day-to-day currency fluctuations by using the average daily spot rate during the quarter of the UBS survey. We use the machinist wage as the most closely comparable wage to German manufacturing wages (and to median OWW wages). We convert UBS wages into EUR at the average USD/EUR exchange rate during the survey quarter and deflate figures with the German CPI (standardized to unity in December 1998).

We also use OWW monthly average wage rates of male workers at the country level for 161 occupations in 155 countries between 1983 and 1999. Missing observations, however, reduce the overlap with MIDI data below the overlap that UNIDO (or UBS) wage data provide. We follow Freeman and Oostendorp's (2001) recommendation and pick the base calibration with lexicographic weighting for the aggregate wages by country. We fill missing values, by country and occupation group, with information from the latest preceding year that has wage information available and reuse OWW wages from 1999 in 2000 and 2001. To mitigate workforce composition effects, we take country medians over 161 OWW occupation groups for foreign wages. We multiply the resulting monthly median occupation wage by twelve to approximate annual earnings for cost function estimation. Complementing foreign OWW wages, we use the German annual earnings survey (table 62321 from *destatis.de/genesis*) and obtain sectoral monthly wages, broken down into three blue-collar and four white-collar occupation groups by sector (two-digit NACE 1.1). We compute median wages over these seven occupation groups by sector and deflate figures with the German

CPI (standardized to unity in December 1998). Occupational wage information from the German annual earnings survey enters the ILO database, on which OWW wages are based, so that these foreign and domestic wages are compatible.

F.3 Complementary data

National accounts information for host-country regressors comes from the World Bank's World Development Indicators and the IMF's International Financial Statistics. To condition selection estimation on skill endowments beyond labor costs, we include the host country's percentage of highly educated residents in 1999 from Barro and Lee (2001) and interact the variable with an indicator whether the percentage exceeds that in Germany (19.5%). We construct market access measures following Redding and Venables (2004), using their measure MA(3). To capture relevant cross-sectional variation, we compute competitor-level averages of the host-country characteristics MNE by MNE. Many host-country regressors are nevertheless statistically insignificant predictors in binary choice estimation, conditional on parent-level observable variables and host-country wages.

G Translog Estimates and Alternative Location Definitions

Table G.1 presents estimates of translog cost function equations for 1,654 stacked MNE observations between 1998 and 2001, as discussed in Subsection III-C. Beyond wages, specifications include turnover and fixed asset regressors, the scaled equivalent of the constant, and indicators of absence from all other locations. Estimates in the upper panel of Table G.1 include the predicted selectivity hazards (inverses of Mills ratios) by location (Assumption A). The lower panel presents estimates from nonparametric selectivity correction (Assumption B), using third-order polynomials in the predicted propensity scores for all locations.

Our definition of aggregate locations is motivated by geographical proximity and broad institutional similarity (Table G.2). As a robustness check, we split the world into the home country and four artificial regions defined by the quartiles of UNIDO manufacturing wages in the initial sample year 1996. Table G.3 reports estimated cross-wage elasticities for the wage-quartile groups of countries, as discussed in Subsection III-F.

Table G.1: TRANSLOG COST PARAMETER ESTIMATES

Employment in:[a]	CEE	DEV	OIN	WEU
	(1)	(2)	(3)	(4)
	Parametric Selectivity Correction (Assumption A)			
ln *Wages*[a]				
HOM	.001	-.013	.027	.054
	(.0006)	(.001)***	(.008)***	(.006)***
CEE	.001	-.008	.008	-.002
	(.0005)**	(.0001)***	(.00004)***	(.00006)***
DEV	-.008	.011	.009	.0006
	(.0003)***	(.001)***	(.0001)***	(.0001)***
OIN	.008	.009	-.086	.043
	(.0004)***	(.0008)***	(.008)***	(.002)***
WEU	-.002	.0006	.043	-.095
	(.0005)***	(.0006)	(.001)***	(.006)***
Selectivity hazard	12.058	24.432	-19.821	35.824
	(11.923)	(13.443)*	(11.606)*	(14.625)**
R^2	.977	.975	.969	.948
	Nonparametric Selectivity Correction (Assumption B)			
ln *Wages*[a]				
HOM	.001	-.008	.023	.059
	(.0006)**	(.001)***	(.007)***	(.006)***
CEE	-.0008	-.006	.007	-.002
	(.0005)	(.0003)***	(.0004)***	(.0005)***
DEV	-.006	.010	.007	-.004
	(.0003)***	(.001)***	(.0007)***	(.0007)***
OIN	.007	.007	-.079	.042
	(.0004)***	(.0007)***	(.008)***	(.002)***
WEU	-.002	-.004	.042	-.096
	(.0005)***	(.0007)***	(.002)***	(.005)***
Series terms				
χ^2 tests (p-value)	495.52 (.000)	246.04 (.000)	151.17 (.000)	244.62 (.000)
R^2	.979	.977	.974	.959

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages).
Notes: Stacked observations of 1,654 MNEs. Further regressors: ln Turnover, ln Fixed assets, ln MNE wage residuals, Absence indicators, Transformed constant (in parametric selectivity regression). Standard errors in parentheses: * significance at ten, ** five, *** one percent. Standard errors corrected for first-stage estimation of selectivity hazards (hence not symmetric on restricted coefficients). Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

[a]Transformed wage-bill shares and regressors.

Table G.2: Aggregate Locations

Locations	Countries
WEU	Western European countries
	(EU 15 plus Norway and Switzerland)
OIN	Overseas Industrialized countries
	including Australia, Canada, Japan, New Zealand, USA
	as well as Iceland and Greenland
CEE	Central and Eastern European countries
	including accession countries and candidates for EU membership
	as well as Balkan countries, Belarus, Turkey, and Ukraine
DEV	Developing countries
	including Russia and Central Asian economies
	as well as dominions of Western European countries and
	of the USA

Table G.3: Cross-wage Elasticities Between Wage Quartile Groups

Employment change (%) in		Wage change (by 1%) in				
		HOM	Qrtl. 4	Qrtl. 3	Qrtl. 2	Qrtl. 1
		(1)	(2)	(3)	(4)	(5)
HOM	*intensive*	-.467**	.402**	.043*	.015*	-.001
Qrtl. 4	*intensive only*	1.193**	-1.339**	.104***	.025*	.009**
	extensive only	.703***	-.763***	.030***	.019***	.004**
Qrtl. 3	*intensive only*	1.026*	.833***	-1.695***	-.190***	.018
	extensive only	.703***	.237***	-.970***	.019***	.004**
Qrtl. 2	*intensive only*	.572*	.317*	-.297***	-.619**	.020
	extensive only	.703***	.237***	.030***	-.981***	.004**
Qrtl. 1	*intensive only*	-.175	.561*	.134	.096	-.624
	extensive only	.703***	.237***	.030***	.019***	-.996***

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages).
Notes: Elasticities at the extensive and intensive margins from 663 stacked MNE observations. Underlying labor demand estimates from parametric selectivity-corrected ISUR estimates (Assumption A). Standard errors from 200 bootstraps: * significance at ten, ** five, *** one percent. Locations: HOM (Germany) and four foreign-country groups by manufacturing-wage quartiles, fourth quartile with top wages.