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ABSTRACT

Employment at multinational enterprises (MNEs) 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. We document a complementarity bias when the extensive margin is excluded and detect salient labor substitution at both margins for German manufacturing MNEs. With a one-percent increase in home wages, for instance, MNEs add 2,000 jobs in Eastern Europe at the extensive margin and 4,000 jobs overall; a converse one-percent drop in Eastern European wages removes 730 German MNE jobs.

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1 Introduction

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. In most locations, the extensive-margin response to permanent wage differentials is about as large as the intensive-margin adjustment and has considerable consequences for overall factor demand. Using comprehensive data on German manufacturing MNEs and their majority-owned foreign manufacturing affiliates, we document that not accounting for extensive-margin adjustment biases conventional intensive-margin estimates predominantly towards complementarity so that common estimates are small or their signs reversed. An instance of complementarity bias arises if firms with a high likelihood to set up shop in low-wage locations also command comparatively low home wages, for example because their relocation propensity serves as a credible threat in wage bargaining at home.

We present a generic MNE model that encompasses motives for both vertical and horizontal foreign direct investment (FDI) and derive a simple estimation framework that accounts for the MNE's regional configuration at both margins. The model offers an integration of two strands of the empirical literature—one on MNEs' location choices and one on MNE operations across existing locations—into a unified procedure.[3] Our estimators extend the univariate sample selection case to one of multiple selections and outcomes. With an eye on ease of implementation, we present conditions under which common Heckman (1979) correction and straightforward non-parametric estimation (similar to Das, Newey and Vella 2003) can be applied location by location and integrated into outcome estimation—in our case a seemingly unrelated equation system of

[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).

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

[3] Among the studies in the former literature are Devereux and Griffith (1998) and Head and Mayer (2004) who research MNEs' location choices. Slaughter (2000), Head and Ries (2002) and Hanson, Mataloni and Slaughter (2005), for instance, analyze MNE operations across existing locations in the latter literature.

the MNE's labor demands.[4] Beyond our context, the estimation technique potentially applies to empirical work on extensive margins in international trade models, such as export-market entry, import-market access, or intra-firm trade.

At the extensive margin, several 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, recognize 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 multinomial logit requires mutually exclusive location choices so that 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 choices in several probit and nonparametric selection regressions, while allowing for correlated decisions. Moreover, our model suggests that the wage impact on location selection should be weighted with the impact of location selection on employment to arrive at the wage effect on an MNE's regional configuration. Doing so shows that wage differentials across locations are substantial predictors of labor substitution within MNEs at the extensive margin.

For the intensive margin, Slaughter (2000) reports that operations in low-wage locations have no detectable impact on relative home employment at U.S. MNEs. Similar to the U.S. evidence,

[4]For consumer demand, there are three main estimation procedures to account for the extensive margin (non-zero demand) and the intensive margin (quantity): Amemiya's (1974) censored system, the primal Kuhn-Tucker approach of Wales and Woodland (1983), and the dual shadow-price approach of Lee and Pitt (1986). As Haefen, Phaneuf and Parsons (2004) remark, most empirical models of consumer demand for multiple goods have relied on the discrete-choice random-utility maximization model (i.e. the extensive margin only). Consumer-demand studies that account for both interior and corner solutions (intensive and extensive margins) have focused on computationally feasible estimation techniques that circumvent the curse of dimensionality associated with primal and dual approaches (Meyerhoefer, Ranney and Sahn 2005). Estimators on the basis of Amemiya (1974) are common. Yen and Lin (2006) generalize Heckman (1979) to a maximum likelihood (ML) estimator for censored demand systems under normality. We present a plausible set of conditions under which an extension of the Heckman two-step estimator is justified instead of ML, and cannot reject the conditions in our MNE sample. We also present assumptions that permit a nonparametric estimator that applies Das et al. (2003) and does not require distributional assumptions while accounting for heteroskedastic error components. Helpman, Melitz and Rubinstein (2008) conduct selectivity correction in the context of country-level bilateral trade flows under the assumption of uncorrelated selection choices; they find the selection bias in bilateral trade to be empirically small, in contrast to MNE expansions in the present paper.

[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.

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. In contrast to the MNE evidence, Feenstra and Hanson (1999) attribute about a third of U.S. relative wage changes to cross-border outsourcing at the sector level, within MNEs or across firms. For Japanese MNEs, Head and Mayer (2004) report that foreign employment in low-income countries does relate to more skill-intensive employment at home. Using additional evidence on U.S. MNEs, Harrison and McMillan (2006) show 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 significant intra-firm trade, so that the net employment effect is small at the intensive margin.[6] When an MNE relocates a production stage that is complementary at the intensive margin, however, job loss can still result at the extensive margin.[7]

We find cross-wage elasticities at both margins to be strictly positive. So, home and foreign employment are substitutes within MNEs not only at the intensive but also at the extensive margin. Bootstrapped standard errors reject equality between the intensive and the total elasticity of substitution for most locations, corroborating the importance of the extensive margin. For overseas developing countries, elasticities are significantly different from zero only at 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.

A third literature on MNEs compares MNE performance to that of other MNEs or national firms with no foreign affiliates. Studies typically 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 (Debaere, Lee and Lee 2006). Using foreign growth rates as instruments for FDI, Desai, Foley and Hines (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. But those studies do not discern whether foreign MNE expan-

[6]Riker and Brainard (1997) report too that affiliate activities in low-income countries are complementary to activities in high-income countries. Hanson et al. (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.

[7]Consistent with this idea, we find the developing-country wage elasticity of home employment not to be significant at the intensive margin but significantly positive at the extensive margin.

sions stabilize industry activity at home or whether market-share gains by MNEs result in market-share losses at domestic competitors. In the interest of concision, we bridge the two literatures on location-selection and employment substitution but do not attempt to integrate the third literature on reallocations between MNEs and non-MNEs. Concretely, we condition on product-market shares as is required for labor-demand estimation within MNEs.

We evaluate the counterfactual question as to how many jobs MNEs would reallocate in response to shrinking wage differentials. A one-percent drop in German wages relative to the sample-mean level would reduce MNE employment in Central and Eastern Europe (CEE) by around 4,000 jobs overall, for instance. Similarly, a one-percent increase in CEE wages would bring 730 jobs to Germany. These are sizeable figures. Wages in CEE are, on average, about 10 percent of the German level in 2000. If the estimated elasticities of substitution were constant at all wage levels, an increase in CEE wages by 450% to cut the wage gap to Germany in half would bring 330,000 ($= 730 \cdot 450$) counterfactual manufacturing jobs to Germany—about a quarter of the estimated home employment at German manufacturing MNEs.[8] Of course, elasticities of substitution are not constant at all wage levels so that the counterfactual prediction is crude. We nevertheless view the magnitude as indicative of the potential importance of multinational labor substitution.

This paper has four more sections. In Section 2, 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 3 discusses the data and descriptive statistics on location choice (details in the Appendix). Estimation results on multinational labor substitution are presented in Section 4, and interpreted in counterfactual evaluations. Section 5 concludes.

2 Multinational Expansion and Operation

2.1 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

[8] 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).

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.[9]

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

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

by Shepard's lemma (see Appendix A).[10]

Not all firms are producing in all locations. The employment effect of MNE selection into locations is both of economic interest in itself and of empirical concern for estimating (1). Consider, for instance, the effect of home wages ($n =$ HOM) on employment in Central and Eastern Europe ($\ell =$ CEE). In the absence of any 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

[9] We adopt the Brown and Christensen (1981, eq. 10.21) short-run version of the Christensen, Jorgenson and Lau (1973) translog cost function and extend the specification to multiple products (Appendix A). A main alternative would be Hall's (1973) generalization of the Diewert (1971) Leontief cost function to the multiproduct case. We favor the translog cost function because it is parsimonious. 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.

[10] 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.

substitutability bias is particularly plausible if institutional uncertainty in the host location, and 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 neighboring WEU countries, where real wages are below German levels between 1996 and 2001, command comparatively low home wages because their relocation propensity serves as a credible threat in wage bargaining at home. For such firms, the uncorrected estimate of the WEU wage-bill response to home wages is negatively biased so that the estimated cross-wage elasticity is biased towards complementarity between home and WEU countries, unless selectivity is controlled for. Previous empirical research largely ignored the selection issue in estimating multinational labor demand.

More formally, an MNE's choice of foreign activity is a two-stage decision problem. 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 $\mathrm{Pr}^{\ell}_{jt}(\mathbf{z}_{j,t-\tau}) = \mathbb{E}\left[d^{\ell}_{jt}|\,\mathbf{z}_{j,t-\tau}\right]$ 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 of cross-location employment responses, the extensive margin cannot be represented with just a count of affiliates or employments because the opening of affiliates has an unobserved effect on MNE employment elsewhere.

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 $y^{\text{ext},\ell}_{jt} \equiv m^{\ell}(\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau}))$.

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, eq. (2) shows that β^{ℓ} coefficients at the intensive margin may be biased unless the unobserved error component $m^{\ell}(\cdot)$ is controlled for. A preview of results in our data documents bias (Table F.1 in the Appendix). Uncorrected cross-wage elasticities are mostly distorted towards complementarity (negative relative differences in the table) and 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 26 percent for DEV wage changes and over-estimated with a substitutability bias of up to .6 percent for CEE wage changes in data on manufacturing affiliates.

Second, eq. (3) shows that estimates of propensity changes in response to wage levels measure only a part of the extensive margin's importance for employment outcomes. Similar to earlier findings, our data exhibit only a weak association between wage levels and location selection, that is $\partial \Pr^{\ell}_{jt}/\partial w^{n}_{t-\tau}$ 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 $\partial y^{\text{ext},\ell}_{jt}/\partial \Pr^{\ell}_{jt}$, wage changes at the extensive margin are found to have an economically and statistically highly significant impact that is about as large as intensive-margin adjustment.

2.2 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 derive the constant-output cross-wage elasticity of substitution between factors ℓ and n.[11] The *cross-wage elasticity of substitution* is defined as $\varepsilon_{\ell n} \equiv \partial \ln y^{\ell}_{jt} / \partial \ln w^n$ and becomes

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

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

$$\frac{\partial s^{\ell}_{jt}}{\partial \ln w^n} = \delta_{\ell n} + \frac{\partial \mathbb{E}\left[\epsilon^{\ell}_{jt} \,|\, \cdot\right]}{\partial w^n_{t-\tau}} \frac{w^{\ell}_t w^n_t}{C_{jt}} = \delta_{\ell n} + \frac{\partial m^{\ell}(\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau}))}{\partial \mathrm{Pr}^{\ell}_{jt}} \frac{\partial \mathrm{Pr}^{\ell}_{jt}}{\partial w^n_{t-\tau}} \frac{w^{\ell}_t w^n_t}{C_{jt}}. \tag{5}$$

The first term in (5) captures the labor demand response at the intensive margin, where $\delta_{\ell n}$ is the regression coefficient on the transformed wage in eq. (1). The second term in (5) is a measure of the labor demand response to a permanent wage change at the extensive margin. The extensive-margin estimate is multiplied by the spot wage w^n_t because estimation on the first stage uses w^n_t as regressors, not their logs. Division by C_{jt}/w^{ℓ}_t converts the extensive-margin estimate from the transformed labor demand eq. (1) back into the wage bill share equivalents.

We presented the derivation of cross-wage elasticities in the benchmark context of competitive labor markets. MNEs, however, are known to pay wage premia over local competitors. Suggested reasons include relatively skilled workforces and theories of rent sharing through efficiency wages or bargaining. Our derived cross-wage elasticities are consistent with departures from competitive labor markets under wage bargaining. 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 reservation wages.[12]

[11] 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.

[12] The first-order condition in Stole and Zwiebel (1996a, 1996b) for single-product firms requires that, at the optimal

The consequences of wage bargaining for labor demand are theoretically ambiguous when contracts are non-binding. A wage-bargaining firm in the original Stole and Zwiebel framework over-employs workers, compared to a neoclassical firm, because additional workers depress other workers' wages by reducing the impact of threats to quit. In a dynamic extension to optimal employment choice over time (Wolinsky 2000), employment at wage-bargaining firms is nearly efficient in the profit maximizing equilibrium. If there is an outside pool of ready-to-be-employed workers, then a wage-bargaining firm under-employs workers in static optimum because replacement workers can be hired instantaneously, lessening the impact of other workers' threats to quit (de Fontenay and Gans 2003). To capture potential employment distortions, whatever their direction, we conduct robustness checks of the competitive labor-market benchmark and augment labor demand equations with industry- and location-specific log wage premia at MNEs.

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. This suggests an extended approach with endogenous output for future research. A structural approach to market-share reallocations after FDI, however, requires assumptions on product-market competition. In contrast, cost function estimation as in (2) is consistent with alternative forms of product-market competition and elucidates employment reallocations within MNEs under lean assumptions. Naturally, the within-MNE employment reallocations documented in this paper are a key part of the labor-market response to FDI also in general equilibrium.

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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.

2.3 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.[13]

Under a parametric specification of the disturbance, we can estimate the sunk costs of entry and exit in probability terms. Given fixed entry costs γ_N^{ℓ} and fixed exit costs γ_X^{ℓ}, the costs of changing presence at location ℓ are

$$G^{\ell}(d_{jt}^{\ell}, d_{j,t-\tau}^{\ell}) = \gamma_N^{\ell}\, d_{jt}^{\ell}(1 - d_{j,t-\tau}^{\ell}) + \gamma_X^{\ell}\,(1 - d_{jt}^{\ell}) d_{j,t-\tau}^{\ell}. \tag{7}$$

So, the decision-relevant sunk cost of presence at ℓ is a function of past presence:

$$F_{j,t-\tau}^{\ell} \equiv G^{\ell}(1, d_{j,t-\tau}^{\ell}) - G^{\ell}(0, d_{j,t-\tau}^{\ell}) = \gamma_N^{\ell} - (\gamma_X^{\ell} + \gamma_N^{\ell})\, d_{j,t-\tau}^{\ell}, \tag{8}$$

where $(\gamma_X^{\ell} + \gamma_N^{\ell})$ is also called the *hysteresis band*. It reflects the sunk costs that induce firms with a presence to continue operations at location ℓ (Dixit 1989).

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

[13] Rearrangement of $H(\mathbf{z}_{j,t-\tau}) \equiv \mathbb{E}_{j,t-\tau}[\Pi(q_{jt}^{\ell,*};\cdot) - \Pi(q_{jt}^{\ell}{=}0;\cdot)] - F_{j,t-\tau}^{\ell}$ shows that

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

where $\Pi(q_{jt}^{\ell,*};\cdot) \equiv p^{\ell} q_{jt}^{\ell,*} - C(q_{jt}^{\ell,*};\cdot)$. This general selection condition encompasses motives for both horizontal and vertical FDI, which may overlap in practice (Feinberg and Keane 2006).

home with $\ell = 1$, the estimation model is therefore

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

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 $\eta^\ell_{j,t-\tau}$ and ϵ^ℓ_{jt}. We consider two sets of assumptions: (1) a parametric version with linear $H(\cdot)$ and joint normality of $\eta^\ell_{j,t-\tau}$ and ϵ^ℓ_{jt}; (2) a nonparametric version for some smooth function $H(\cdot)$ with independent $\eta^n_{j,t-\tau}$ and ϵ^ℓ_{jt} for $n \neq \ell$.

Assumption 1: 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 (1). 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. 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.

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.

[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.

Assumption 2: 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 applies a related result from Das et al. (2003).

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 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 1, 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.

3 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.[15] We restrict the sample to manufacturing parents and

[15]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

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.[16]

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 E.

3.1 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 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 corrected for within-MNE shipments, but is a proxy nevertheless to affiliate production for cost-function estimation.

For German parent firms, employment, turnover and fixed-asset information comes 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 popu-

of the German MNEs in our sample are affiliates of foreign MNEs.

[16]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 CEE 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 prevalent in the sample with affiliates from any sector (lower panel of Table F.1 in the Appendix).

larity 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.[17] 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 F.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 nevertheless conceal considerable heterogeneity so that for robustness estimation we also group 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-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.[18] 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

[17]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.

[18]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 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	.0002	.00007	.0002
All German MNEs' total employment share	.175[a]	.014	.002	.006	.021
MNE log wage premia over local competitors[b]		.626	.861	.072	.283

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).

totals. Although German manufacturing MNEs employ an estimated 17.5 percent of German manufacturing workers, the MNEs' labor-market power is arguably limited. In Germany, collective agreements with strong industry-wide unions do not allow individual firms to deviate from wage schedules that specify wages by worker skill and seniority (only distressed firms qualify for exception clauses). German 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). Wage premia also 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). However, to control for potential labor-demand distortions from MNE rent sharing, we will account for MNE wage premia by host country and industry below.

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-

Table 2: LOCATION COUNTS BY MNE

L in 1996	L in 2000 1	2	3	4	5	*Total* (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).

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 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

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.

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.[19] Motivated by these findings, we define the

[19]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

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.

3.2 Wage data

Paid wages are not reported in MIDI. We use manufacturing wages by country and sector for 1996 through 2001 from the UNIDO Industrial Statistics Database at the 3-digit ISIC level (dividing sectoral wage bills by employment) for our main analysis. We also report robustness checks using OWW wage data by occupation (Occupational Wages around the World, Freeman and Oostendorp 2001).[20] Appendix E.2 provides details on the wage data sources. Though USTAN has German parent wages, for comparability we take German wages from the same outside sources as affiliate wages.

We construct different wage regressors for the two margins to address econometric concerns. For intensive-margin labor-demand estimation on the second stage, we use median wages over sectors by country. The median mitigates possible sectoral workforce composition effects behind local wages. Concretely, we take the arithmetic mean over the sector-median wages across the foreign countries where the MNE is present in a given year, and we take Germany-wide sector medians of the home wages by year. These wages are the decision-relevant local labor costs that the MNE faces at the intensive margin. To account for typical MNE wage premia on top of local labor costs, we obtain Swedish affiliate wages by host country and sector 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 sector in 1998, and use the log ratios as controls for wage premia in robustness checks. German and Swedish MNEs exhibit similar labor-demand behavior abroad (Becker et al. 2005).

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 (sector-median wages by country) MNE-specific. We take competitor averages for every MNE over the foreign wages that the MNE's Ger-

net changes are mostly country changes and not reentries with different affiliates.

[20] We report additional robustness checks with UBS wage data in the working paper version. Results vary little with different wage data.

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	.379	.323	.299	.702
Indic.: Presence in $t-\tau$	1	.351	.296	.281	.706
MNE-wide regressors (Labor-demand estimation)					
Wage bill share (t)	.791	.067	.049	.170	.191
ln Turnover (t)	18.450	15.931	16.505	17.277	17.073
ln Fixed assets (t)	17.264	14.886	15.108	15.804	15.282
ln Wage (t)	10.360	8.286	8.657	10.316	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 $(t-\tau)$	11.234	10.525	12.637	12.826	11.552
Comp.s' hosts skill share $<$ Home $(t-\tau)$	20.151	18.958	22.358	22.565	20.715
Comp.s' hosts skill share $\geq$ Home $(t-\tau)$	42.100	39.052	48.083	49.629	43.382
Comp.s' hosts distance $(t-\tau)$	31.669	29.505	35.930	36.562	32.620
Comp.s' hosts ln Cons. p.c. $(t-\tau)$	30.444	28.614	34.007	34.534	31.243
Parent-firm regressors (Selection estimation)					
Indic.: Headquarters West Germany $(t-\tau)$	.973	.964	.974	.969	.974
ln Count of host countries $(t-\tau)$	1.138	1.327	1.638	1.478	1.263
ln Employment $(t-\tau)$	6.342	6.452	7.214	6.880	6.474
ln Equity $(t-\tau)$	16.662	16.852	17.837	17.588	16.941
ln Liability $(t-\tau)$	17.728	17.927	18.716	18.373	17.891
ln Capital-labor ratio $(t-\tau)$	10.835	11.004	11.070	11.104	10.936
Parent observations	1,640	612	457	489	1,095

Sources: MIDI and USTAN 1996 to 2001, censored (second-stage) estimation sample of 1,640 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).

man competitors pay.[21] The wage in CEE, for example, is the average wage in the CEE countries where competitors' affiliates were located. Competitors' labor costs abroad are arguably among the important decision variables for an MNE's location choice. We apply the same procedure to all other host-country characteristics. For the annual home wage, we use sector-mean wages because wage variables that reflect the workforce composition are valid predictors at the extensive margin.

3.3 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 our main estimation specification, we consider multinational labor demand during the years 1998-2001 for a sample of 1,640 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 322 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 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

[21] We consider only competitors within an MNE's broad manufacturing sector. The eight sectors 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 regressor 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 $t - \tau$ ($\tau = 2$).

[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 are less sensitive to missing information.

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. For all other parent-firm regressors, regional conditional means (columns 2 to 5) exceed the unconditional mean (column 1), and regional means tend to be the lower the higher the frequency of MNE presence. Conditional on their presence abroad, MNEs exhibit larger home workforces, larger parent-firm equity or debt, and higher parent-firm capital-labor ratios.

4 Estimation

We estimate the MNE model—selection equations (6) and outcomes (2)—for all locations under Assumptions 1 and 2. Exclusion restrictions and timing provide identification.

4.1 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 consider German home wages as exogenous to the individual MNE because German manufacturing firms face bargained wage schedules from industry-specific collective agreements between employer associations and strong unions. The threat of employment relocation abroad arguably 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 German home wages are adequately identified at both margins from cross-sectoral variation. Time variation in home wages provides additional identification.

We measure foreign labor costs with location-wide sector medians of wages and consider those costs as 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. 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.

Wage premia at MNEs, however, are a sign of departure from competitive labor markets and consistent with rent sharing between the firm and its workers. As shown above, the translog cost specification is consistent with wage-bargaining under non-binding contracts if we use reservation-wage measures in estimation, such as location-wide median wages. A remaining empirical concern is that affiliates' paid wages might bias the reservation-wage coefficient when paid wages are omitted from labor-demand estimation on the second stage. Suppose the disturbance includes an MNE-wage premium over local reservation wages. A particular concern for our argument is then that the omitted error component would bias cross-wage elasticities by distorting the wage coefficient in the outcome equation. To check for the robustness of our results to such distortion, we use industry- and location-specific wage premia at Swedish MNE affiliates and include them as regressors.

Serial correlation in the selection disturbance, due to persistence in unobserved local demand 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 estimates of labor demand are based on constant-output cost-function estimation so that 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, financial measures and productivity-related variables such as profits per equity. On the second stage, inverse Mills ratios or nonparametric propensities of foreign presence control for heterogeneity and the MNEs' motives to conduct FDI. A remaining unobserved MNE-specific performance advantage, such as global productivity say, 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.

4.2 Location choice

We first estimate location-selection equations (6)

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

Probit estimation. For probit estimation, we start by investigating the implication of Assumption 1 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).

[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 or future output, however, we find broadly similar cross-wage elasticity estimates at both margins (not reported for brevity) 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.

Table 5: Marginal Effects in Pooled Probit Regressions

Predictors $(t-2)$ / Presence (t)	CEE (1)	DEV (2)	OIN (3)	WEU (4)
FDI in CEE	.619	.184	.472	-.361
	(.234)***	(.270)	(.299)	(.293)
FDI in DEV	-.001	.800	-.094	-.054
	(.109)	(.111)***	(.070)	(.149)
FDI in OIN	-.259	-.485	-.083	-.179
	(.476)	(.326)	(.442)	(1.035)
FDI in WEU	.314	.108	.009	.983
	(.203)	(.297)	(.298)	(.019)***
Home sector wage	.0004	.001	.006	.019
	(.004)	(.004)	(.003)*	(.007)**
Competitors' wages CEE	-.050	-.023	.001	-.099
	(.055)	(.045)	(.039)	(.060)*
Competitors' wages OIN	-.001	-.002	-.028	.025
	(.015)	(.016)	(.015)*	(.020)
FDI in loc. $\times$ Home sector wage	-.0007	-.005	-.015	-.020
	(.005)	(.004)	(.004)***	(.008)***
FDI in CEE $\times$ Comp.s' wages CEE	.054	-.060	-.093	.090
	(.066)	(.057)	(.050)*	(.083)
FDI in OIN $\times$ Comp.s' wages OIN	.010	.029	.035	.005
	(.027)	(.026)	(.019)*	(.034)
$\ln$ Count of host countries	.036	.086	.031	.128
	(.040)	(.035)**	(.028)	(.053)**
$\ln$ Employment	.116	.057	.064	.153
	(.026)***	(.023)**	(.021)***	(.031)***
$\ln$ Liability	-.089	-.047	-.052	-.166
	(.022)***	(.019)**	(.017)***	(.026)***
$\ln$ Capital-labor ratio	.085	.023	.034	.072
	(.022)***	(.019)	(.017)*	(.026)***
Obs.	2,413	2,413	2,413	2,413
Pseudo R^2	.559	.523	.555	.457

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 in DEV and WEU and their interactions with FDI presence, Competitors' host-country $\ln$ Market access, Indic. of Headq. West Germany, $\ln$ Equity, Parent profits/equity, Competitors' host-country skill shares, Competitors' host-country distance, Competitors' host-country $\ln$ Consumption per capita. Without wage-presence interactions, past presence has a marginal effect of .779 (std. err. .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 plausibility of Assumption 1 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 4.4.[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.

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. 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 only need home-wage coefficients for the cross elasticities of substitution at the extensive margin (there is no extensive margin for home where foreign wages would enter). Bootstraps over both estimation stages will show even for the statistically weak home-wage prediction of lo-

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

cation selection into CEE that, weighted with the strong labor demand effects of CEE selection, the home wage significantly affects the elasticities of labor substitution at the extensive margin.

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 below .001. German MNEs with large home employment, low parent debt, and a high capital-labor ratio at the parent firm two years ago are significantly 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, parent equity, and parent profits per equity, 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 .01 (DEV) and .64 (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]

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 4.4.

To gain a sense of sunk entry and exit costs behind hysteresis, we run a short descriptive regression of presence in the year-2000 cross section on past presence at the location and any other location in 1996. In this short regression, past presence is a statistically highly significant predictor of presence four years later at the same location; presence elsewhere serves as a rudimentary control and is also highly statistically significant. These descriptive estimates provide an indication of sunk cost components in probability terms. Recall that the sunk cost part of location choice in eq. (8) can be represented as the difference between sunk entry costs and the hysteresis band. Table 6 shows the result of this decomposition based on coefficient estimates for the short descrip-

[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 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: SUNK ENTRY AND EXIT COSTS AT FOUR-YEAR HORIZON

Current presence (2000)	CEE	DEV	OIN	WEU
	(1)	(2)	(3)	(4)
Sunk entry cost: γ_N (1996)	.525***	1.069***	1.156***	.441***
Sunk exit cost: γ_X (1996)	.902***	.412***	.558***	.668***
Hysteresis band: $\gamma_N + \gamma_X$ (1996)	1.427***	1.481***	1.714***	1.109***
Marginal effect of hysteresis band (1996)	.518***	.512***	.561***	.421***

Sources: MIDI 1996 and 2000, 867 manufacturing MNEs and their majority-owned foreign manufacturing affiliates.
Notes: Estimates are probit coefficients from a descriptive regression of current presence indicators at a location on past presence indicators at the location and any other location. Significance levels from χ^2 tests: * significance at ten, ** five, *** one percent. . Foreign locations: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

tive regression.[27] Past presence in a given location increases the likelihood of presence four years later by fifty percent in all but WEU, where the marginal effect predicts a more than forty percent increase. These estimates are lower than the around seventy-percent predictions at the two-year horizon (see above), but still substantial. The descriptive decomposition suggests that entry costs are largest in the distant low-income and high-income locations DEV and OIN, and dominate exit costs there. Conversely, entry costs are lowest in the nearby low-income and high-income locations CEE and WEU, and significantly smaller than exit costs. Among the exit costs are the opportunity costs of absence. German MNEs are considerably less reluctant to leave distant locations DEV and OIN than they abandon the neighboring locations CEE or WEU.

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

[27]The sunk cost decomposition involves an estimate of the constant so that entry and exit costs cannot be expressed in marginal probability terms of their own. A marginal probability measure can be inferred for their sum, the hysteresis band.

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 7 reports coefficient estimates by location. The predicted propensity scores of location choice are .338 for CEE, .291 for DEV, .262 for OIN and .617 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.

Table 7 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.

4.3 Labor demand estimation with selectivity correction

We proceed to estimate outcomes (2)

$$y^{\ell}_{jt} = \mathbf{x}^{\ell}_{jt}\beta^{\ell} + m^{\ell}(\mathbf{P}_{jt}(\mathbf{z}_{j,t-\tau})) + \epsilon^{\ell}_{jt}$$

Table 7: MARGINAL EFFECTS IN NONPARAMETRIC PROBABILITY MODEL

Predictors ($t-2$) / Presence (t)	CEE (1)	DEV (2)	OIN (3)	WEU (4)
FDI in CEE	.644	.108	.193	-.207
	(.145)***	(.149)	(.138)	(.184)
FDI in DEV	-.070	.383	-.065	-.007
	(.088)	(.116)***	(.083)	(.107)
FDI in OIN	.016	.060	.068	.075
	(.553)	(.568)	(.550)	(.687)
FDI in WEU	.174	-.122	-.057	1.082
	(.222)	(.215)	(.201)	(.258)***
FDI[a] in loc. $\times$ Home sector wage	.001	.006	-.010	-.004
	(.003)	(.004)*	(.003)***	(.004)
FDI in OIN $\times$ Comp.s' wages OIN	-.001	-.002	.031	-.003
	(.018)	(.018)	(.017)*	(.022)
Series terms of wages: p-values from F tests				
Home sector wage terms		.041	.021	
Competitors' CEE wage terms				
Competitors' DEV wage terms				
Competitors' OIN wage terms	.012	.052		
Competitors' WEU wage terms				
$\ln$ Employment	.064	.039	.049	.090
	(.014)***	(.014)***	(.013)***	(.017)***
$\ln$ Liability	-.046	-.028	-.036	-.094
	(.011)***	(.012)**	(.011)***	(.014)***
$\ln$ Capital-labor ratio	.046	.020	.028	.045
	(.011)***	(.012)*	(.011)***	(.014)***
Obs.	2,413	2,413	2,413	2,413
R^2	.666	.618	.633	.556

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): Interactions of competitors' wages in CEE/DEV/WEU with FDI presence in CEE/DEV/WEU, Competitors' host-country $\ln$ Market access, $\ln$ Count of host countries, Indic. of Headquarters West Germany, $\ln$ Equity, Parent profits/equity, Competitors' host-country skill shares, Competitors' host-country distance, Competitors' host-country $\ln$ Cons. p.c. 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).

[a]FDI presence in regression location.

for all locations. 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). For parametric correction, we include the predicted selectivity hazard (inverse of the Mills ratio) from the first stage among the regressors. Under nonparametric correction, we include the predicted propensity scores from the first-stage estimates. We use cross validation to choose the order of polynomial expansion at this stage. A third-order approximation performs better in CEE and DEV (but worse in OIN and WEU); we use a third-order approximation because much of our interest lies on CEE. 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). We treat induced heteroskedasticity following Heckman (1979), resulting in different standard errors on symmetric coefficients.

Table 8 presents estimates of translog cost function equations for 1,640 stacked MNE observations between 1998 and 2001. Beyond the reported wage coefficients, the equations include the full sets of turnover and fixed asset regressors, the scaled equivalent of the constant, and indicators of absence from all other locations. All but two wage coefficients in Table 8 are significantly different from zero at the one percent level, and all coefficients but one are significant at the five percent level in each, parametric and nonparametric, regression. Most coefficients on output and fixed assets (not reported) are similarly highly significant.

Estimates in the upper panel of Table 8 include the predicted selectivity hazards (inverses of Mills ratios) by location (Assumption 1). Selectivity hazards are statistically different from zero at the one percent level in all equations except DEV (significance at ten-percent level). The lower panel presents estimates from nonparametric selectivity correction (Assumption 2), using third-order polynomials in the location's propensity score interacted with indicators for presence at all other locations. χ^2 tests on the series terms overwhelmingly reject their joint equality to zero. The

Table 8: TRANSLOG COST PARAMETER ESTIMATES

Employment in:[a]	CEE	DEV	OIN	WEU
	(1)	(2)	(3)	(4)
Parametric Selectivity Correction (Assumption 1)				
ln *Wages*[a]				
HOM	.020 (.001)***	-.002 (.0008)**	.078 (.004)***	.183 (.005)***
CEE	-.008 (.0008)***	-.001 (.0002)***	-.003 (.0004)***	-.008 (.0005)***
DEV	-.001 (.0003)***	.001 (.0008)	-.002 (.0004)***	.004 (.0006)***
OIN	-.003 (.00007)***	-.002 (.00007)***	-.112 (.003)***	.039 (.001)***
WEU	-.008 (.0001)***	.004 (.0001)***	.039 (.001)***	-.219 (.004)***
Selectivity hazard	81.487 (15.830)***	32.872 (17.751)*	33.468 (12.462)***	92.618 (16.618)***
R^2	.945	.950	.966	.932
Nonparametric Selectivity Correction (Assumption 2)				
ln *Wages*[a]				
HOM	.022 (.001)***	.001 (.001)	.073 (.005)***	.145 (.006)***
CEE	-.007 (.0008)***	-.003 (.0004)***	-.003 (.0005)***	-.009 (.0006)***
DEV	-.003 (.0004)***	.0008 (.0009)	-.002 (.0006)***	.003 (.0007)***
OIN	-.003 (.0005)***	-.002 (.0006)***	-.109 (.005)***	.040 (.002)***
WEU	-.009 (.0006)***	.003 (.0007)***	.040 (.002)***	-.179 (.006)***
Series terms				
χ^2 tests (p-value)	618.4 (.000)	457.6 (.000)	183.5 (.000)	293.0 (.000)
R^2	.956	.957	.967	.929

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages).
Notes: Stacked observations of 1,640 MNEs. Further regressors: ln Turnover, ln Fixed assets, 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 9: 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*	-.574***	.051***	.011	.150***	.361***
CEE	*intensive only*	1.596***	-1.295***	-.039	-.081	-.181
	extensive only	.795***	-1.250***	.071	.155	-.097
DEV	*intensive only*	.651	-.071	-.912***	-.116	.448**
	extensive only	.772***	-.250	-.982***	.324	.656
OIN	*intensive only*	2.328***	-.040	-.031	-3.160***	.903***
	extensive only	.960***	-.288	.032	-2.597*	.365
WEU	*intensive only*	2.214***	-.036*	.048**	.358***	-2.584***
	extensive only	1.016***	-.341	.128	1.137*	-.951***

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages).
Notes: Elasticities at the extensive and intensive margins from 1,640 stacked MNE observations. Underlying labor demand estimates from parametric selectivity-corrected ISUR estimates (Assumption 1, Tables 5 and 8). Standard errors from 200 bootstraps: ** significance at five, *** one percent. Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

translog cost function regressors predict the bulk of labor demand variation across locations, with the R^2 goodness of fit ranging between .93 and .97 across equations. The regression fit is similar under parametric and nonparametric selectivity correction. Overall, we view the significance of selectivity correction terms as evidence for the importance of the extensive margin.

Elasticities of multinational labor substitution. We repeat joint selection (6) and outcome (2) estimation in 200 bootstraps to infer elasticities of labor substitution (4) and their standard errors at both margins.[28]

Table 9 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 elasticities are not additive by eq. (4).

[28]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 widely differ from bootstrapped confidence intervals.

Own-wage elasticities along the diagonal—for both intensive and extensive margins—are uniformly and significantly 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 is considerably larger in most foreign locations than at home, suggesting that MNE employment abroad responds more sensitively to labor costs there than home employment responds to home wages.

Cross-wage elasticities in the first row (foreign wage effects on home employment) and in the first column (home wage effects on foreign employment) are significantly positive for eleven out of thirteen estimates at the intensive and the extensive margins. A one-percent reduction in the wage in CEE, for instance, is associated with a .05 percent drop in home employment at German MNE parents. In contrast, a one-percent increase in the German sector wage is associated with a 1.6 percent boost to MNE employment in CEE at the intensive margin and a .8 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 a noticeable component of adjustment, beyond its crucial role in correcting cost function estimates for location selectivity bias. We find that elasticities at the extensive margin are strictly positive. So, home and foreign employment are substitutes within MNEs not only at the intensive but also at the extensive margin. Although the CEE and DEV home wage effects on selection were not statistically different from zero on the first stage with probit (Table 5), the strong significance of the selection effect on labor demand on the second stage in CEE (selectivity hazard coefficient in Table 8) turns home wage effects into significant predictors of employment substitution at the extensive margin.

Elasticities at the extensive margin are smaller in magnitude than at the intensive margin in the geographically close locations CEE and WEU, and in OIN. In DEV, however, the extensive margin dominates the statistically insignificant elasticity at the intensive margin and we find a .8 percent increase in DEV employment in response to a one-percent home wage increase—similar in magnitude to that in CEE.

Bootstrapping allows us to test whether the elasticities at the intensive margin are statistically significantly different from the total elasticities. We reject equality for DEV, OIN and WEU (with t statistics between 2.1 and 16.6) on UNIDO wages and reject their equality for all locations (t statistics between 4.1 and 21.4) on OWW wages, corroborating the importance of the extensive margin.

Cross-wage estimates beyond the first row and column are for the most part not statistically different from zero. Notable exceptions at the intensive margin are significant pairs of positive cross-wage effects involving WEU: on the one hand of OIN on WEU (.36) and vice versa (.90), and on the other hand of DEV on WEU (.05) and vice versa (.45). The significantly positive and mutually consistent effects suggest that MNE employment is a substitute at the intensive margin between OIN and WEU and between DEV and WEU. The substitution effect is also corroborated by a positive cross-wage elasticity between OIN and WEU (1.14) at the extensive margin.

Overall, our estimation strategy finds labor in one location to be a substitute to labor elsewhere. In the absence of a correction for the extensive margin, conventional translog estimates for the intensive margin in our data would result in under-estimated coefficients for 15 out of 25 coefficients—a bias towards complementarity (see Table F.1 in the Appendix). 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 and foreign employment to be complements in any industry or location, whereas Harrison and McMillan (2006) find complementarity for wages in low-income locations and home employment. Harrison and McMillan (2006) restrict the sample to manufacturing affiliates, as we do. So, 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.

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. 1, UNIDO 98-01	-.574	.051	.011	.150	.361	1,640
	(.062)***	(.007)***	(.008)	(.028)***	(.037)***	
Ass. 1, UNIDO 00	-.631	.062	.034	.202	.332	322
	(.115)***	(.026)**	(.021)	(.071)***	(.078)***	
Ass. 1 AR(2), UNIDO 98-01	-.576	.051	.012	.151	.363	1,640
	(.081)***	(.014)***	(.012)	(.034)***	(.046)***	
Ass. 1, UNIDO 98-01, IUI Δw	-.592	.048	.011	.168	.365	1,640
	(.056)***	(.007)***	(.007)	(.029)***	(.033)***	
Ass. 1, OWW 98-01	-.477	.051	-.002	.209	.219	1,458
	(.053)***	(.010)***	(.005)	(.030)***	(.037)***	
Ass. 2, UNIDO 98-01	-.525	.053	.015	.144	.313	1,640
	(.051)***	(.007)***	(.007)**	(.024)***	(.035)***	
Omnipresent MNEs						
Ass. 1, UNIDO 98-01	-1.354	.090	-.021	.526	.758	93
	(.209)***	(.104)	(.048)	(.135)***	(.143)***	

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

4.4 Specification comparisons

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).

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 1 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-

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, IUI 98-01 Δw	OWW 98-01	UNIDO 98-01	UNIDO 98-01
Emplmt. chg. (%)	Ass. 1	Ass. 1	Ass. 1	Ass. 1	Ass. 1	Ass. 2	Ass. 1
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CEE	1.596	1.810	1.599	1.503	1.366	1.648	3.535
	(.218)***	(.748)**	(.464)***	(.194)***	(.247)***	(.226)***	(4.062)
DEV	.651	1.534	.669	.650	-.147	.880	-.444
	(.466)	(1.004)	(.722)	(.430)	(.480)	(.397)**	(1.072)
OIN	2.328	2.573	2.330	2.593	3.540	2.235	1.938
	(.432)***	(.888)***	(.572)***	(.421)***	(.516)***	(.363)***	(.482)***
WEU	2.214	1.860	2.222	2.235	2.087	1.915	2.851
	(.224)***	(.407)***	(.287)***	(.201)***	(.353)***	(.205)***	(.494)***
Obs.	1,640	322	1,640	1,640	1,458	1,640	93

Sources: MIDI and USTAN 1996 to 2001 (UNIDO and OWW wages, IUI wage premia).
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 five, *** one percent. Locations: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

wage elasticity estimates. In row four, we include industry-specific MNE wage premia by location on the second stage, using IUI wage data for affiliates of Swedish MNEs for 1998; only minor changes to coefficient estimates result. OWW wage data in the fifth row lead to smaller estimation samples but coefficient estimates are similar across wage data.[29]

Nonparametric estimation under Assumption 2 does not yield statistically different estimates, excepting DEV (row six). The subsample of 93 omnipresent MNEs between 1996 and 2001 is small but results in significant outcome estimates on the second stage (last row); the magnitude of cross-wage elasticity estimates, when significant, is considerably larger than for the stacked samples, suggesting that home employment at omnipresent MNEs responds more elastically to foreign wage changes. Estimates for DEV are not significant except for nonparametric specifications. This is consistent with the assertion that higher-order series terms in the outcome regression help remove bias that parametric selectivity correction cannot prevent with a single selectivity hazard.

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

[29]We find similar results using UBS wage data.

Table 12: HOME-WAGE ELASTICITIES AT THE EXTENSIVE MARGIN

	Home wage change (1%), by regression specification						
	Stacking						Omnipr.
	UNIDO 98-01 Ass. 1	UNIDO 00 Ass. 1	UNIDO 98-01 AR(2) Ass. 1	UNIDO, IUI 98-01 Δw Ass. 1	OWW 98-01 Ass. 1	UNIDO 98-01 Ass. 2	UNIDO 98-01 Ass. 1
Emplmt. chg. (%)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CEE	.795	.838	.911	.769	.395	-.040	.643
	(.201)***	(.232)***	(.289)***	(.058)***	(.380)	(9.586)	(.300)**
DEV	.772	.572	1.200	.603	.975	3.941	.592
	(.162)***	(.252)**	(.580)**	(.182)***	(.298)***	(17.680)	(.503)
OIN	.960	1.116	.774	.792	1.431	-4.249	.345
	(.340)***	(.392)***	(.272)***	(.182)***	(.845)*	(7.373)	(.331)
WEU	1.016	1.183	.521	.979	1.561	-2.457	.719
	(.171)***	(.301)***	(.759)	(.093)***	(.372)***	(3.141)	(.096)***
Obs.	1,640	322	1,640	1,640	1,458	1,640	93

Sources: MIDI and USTAN 1996 to 2001 (UNIDO and OWW wages, IUI wage premia).
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 five, *** one percent. Locations: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

several other specifications. In fact, the comments on the rows of Table 10 above apply also to the columns of Table 11, except only that the significant cross-wage elasticity estimates for the subsample of omnipresent MNEs now closely resemble those from other specifications.

At the extensive margin, Table 12 documents that home-wage elasticities of foreign employment are (highly) significant in the parametric specifications (columns 1 through 5). Neither an AR(2) error specification for selection, nor the inclusion of industry and location-specific MNE wage premia on the second stage, nor the use of OWW wage data yield a significantly different elasticity estimate at any location. Under AR(2) selection disturbances, CEE and DEV point estimates increase while OIN and WEU point estimates drop (the latter to a level that is statistically indistinguishable from zero but also statistically indistinguishable from the benchmark). With OWW wage data, all elasticity point estimates but the one for CEE increase. 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 (columns 5 and 6 of Table 12) are not statistically different from zero but similar in magnitude when plausible (column 5, excepting DEV). 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 seem to provide a precise estimate of the extensive margin itself. The similarity between parametric and plausible nonparametric estimates is nevertheless an indication that our parametric benchmark estimates of the extensive margin are reasonable. Point estimates for omnipresent MNEs (column 7) are smaller than in the benchmark specification, arguably because this selected sample expands to foreign locations more frequently.

In summary, robustness checks confirm the statistical plausibility of the benchmark estimates in Table 9 under parametric selectivity correction (Assumption 1). Nonparametric estimates (Assumption 2) are similar and highly significant at the intensive margin, but fail to attain statistical significance at the extensive margin.

4.5 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 F.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 and do not show statistically significant home-wage elasticities on foreign employment at the selection margin, similar to the distant DEV location before. These two facts corroborate our findings for the aggregate locations. Third, estimates at the selection margin show scant variability off the diagonal for any given column. 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 outcome-margin substitutability cuts across the artificial four-quartile regions more frequently than across the geographically and institutionally related aggregate loca-

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 at the intensive margin on	CEE (1)	DEV (2)	OIN (3)	WEU (4)
Home[a]	728	161	2141	5143
	(101)***	(118)	(401)***	(526)***
Foreign[b] *extensive margin*	-1,954	-2,567	-3,066	-4,010
	(493)***	(537)***	(1084)***	(674)***
Foreign[b] *total*	-3,951	-2,128	-7,999	-9,656
	(734)***	(1698)	(1933)***	(1162)***

Sources: Own calculations based on selectivity corrected translog estimates for 1,640 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 1, Table 9) multiplied by employment in 2000 (Table 1). Standard errors from 200 bootstraps: ** significance at 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).

tions. The latter two facts support our definitions of aggregate locations as more coherent.

4.6 Counterfactual evaluation

We now 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? Counterfactual predictions in Table 13 give answers to these questions.

We use the home-wage elasticities of foreign employment and the foreign-wage elasticities of home employment from benchmark estimates in Table 9. 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.

A one percent smaller wage gap between Germany and locations in CEE, for instance, is associated with around 730 more jobs at German parents and 4,000 less jobs at affiliates in CEE. CEE affiliates tend to have smaller work forces and, arguably, lower labor productivity than German establishments so that employment in CEE is more sensitive to home wage changes than home employment responds to foreign wages. The labor substitution effects of one-percent wage changes between home locations and CEE are smaller than the effects relative to OIN or WEU. In absolute magnitude, however, a closing of the HOM-CEE wage gap by half *at constant elasticities* results in larger employment effects than a reduction of the HOM-OIN or HOM-WEU wage gaps by half. Using country populations as weights for location mean UNIDO wages, CEE wages are, on average, 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 are constant at all levels of wages, 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 330,000 $(= 730 \cdot 450)$ counterfactual manufacturing jobs (with a standard error of 45,000 jobs) to Germany—around a quarter of the estimated home employment at German manufacturing MNEs in 2000 (Table 1). 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). 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 about half would attract only 70,000 counterfactual manufacturing jobs to the German plants of German manufacturing MNEs.

Elasticities of labor substitution are local properties of the MNE's cost function, however, and the assumption of a constant elasticity of substitution at all wage levels is coarse. The rough calculations above are merely intended to put an economic meaning to the abstract elasticity figures. In our view, the magnitude of our calculations for constant elasticities nonetheless underscores the potential importance of job substitution within MNEs for labor market outcomes. The calculations also highlight that marginal elasticity estimates alone can be misleading indicators of job loss unless the elasticity point estimates are weighted with the prevailing wage gaps.

5 Conclusion

The public discourse on offshoring espouses the idea that multinational enterprises (MNEs) substitute jobs at home for foreign employment. But economic studies on MNE labor demand across locations have found weak or no evidence of job substitution. We integrate two distinct branches of the literature—one on predictions of MNEs' location choices, and one on labor substitutability across established MNE locations—into a single econometric model that corrects labor-demand estimation for location selectivity. 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. We derive conditions for common Heckman (1979) selectivity corrections, location by location, and for nonparametric identification. Our novel estimation strategy detects a frequent complementarity bias in estimates from conventional uncorrected methods.

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 across locations. With every percentage increase in Central and Eastern European wages, for instance, German manufacturing MNEs are found to allocate 730 MNE jobs to Germany. Similarly, with every percentage increase in German wages, German MNEs allocate 2,000 jobs to Central and Eastern Europe at the extensive margin and 4,000 jobs in total. Given the sizeable wage differential between Germany and Central and Eastern Europe (requiring a 450 percent increase in Eastern European wages in 2000 to reduce the gap by half), we conclude that international wage differentials have a pronounced impact on multinational labor substitution. As the wage gaps to CEE countries narrow, our estimates lead us to expect that CEE jobs are relocated to Germany. This prediction is consistent with recent industry and press reports on German MNEs that repatriate jobs.

The estimated 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 foreign direct investment might gain product-market shares and consequently employment. But, second, competitors at whose expense MNEs' product-market shares expand might lose employment. Measuring the net employment effect in market equilibrium remains a task for future research. Naturally, an essential aspect of reallocations in equilibrium is the production reallocation within MNEs at the extensive and the intensive margin.

Appendix

A Multiproduct translog cost function

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

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

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 1 *The disturbances* $(\epsilon^{n}_{jt}, \eta^{\ell}_{j,t-\tau})$ *are multivariate normally distributed with* $\epsilon^{n}_{jt} = \lambda\, e_{jt} + \pi^{n}_{\epsilon}\, \upsilon^{n}_{jt}$ *and* $\eta^{\ell}_{j,t-\tau} = \sqrt{1-\omega}\, e_{jt} + \sqrt{\omega}\, u^{\ell}_{jt}$*, where* $\omega \in [0,1]$ *and the standard normal variables* $e_{jt}, u^{\ell}_{jt}, \upsilon^{n}_{jt}$ *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 1 does this. Under Assumption 1, the variances and covariances of the selection shocks are $\sigma^{\ell\ell}_{\eta} = 1$, as is common for probit, and $\sigma^{n\ell}_{\eta} = 1-\omega$. 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 1 is testable. We obtain estimates of $\sigma^{n\ell}_{\eta} = 1-\omega$ 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 parametric 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 1 holds.*

Proof. Denote the standard normal density and distribution functions with $\phi(\cdot)$ and $\Phi(\cdot)$. Under Assumption 1, 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 1), 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 2 formally states one set of sufficient conditions for identification.

Assumption 2

(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 2 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 2) and polynomial series estimation, the derivatives of $m^\ell(\cdot)$ and P^ℓ_{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 Data

E.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

[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.

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.

E.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 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 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 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. 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.

E.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.

F Alternative Estimators, Samples and Definitions

Table F.1 presents the relative difference between conventional estimates of cross-wage elasticities at the intensive margin and selectivity-corrected estimates. The reported numbers are the conventional estimate less the selectivity-corrected estimate, divided by the selectivity-corrected estimate. Uncorrected cross-wage elasticities are frequently distorted towards complementarity (negative relative differences), especially in the important first row and first column of the cross-wage elasticity matrix, and signs are reversed into outright complementarity in several instances (relative differences of less than negative one). Distortion into complementarity is observed especially often in the sample with affiliates from any sector (lower panel of Table F.1).

Our definition of aggregate locations is motivated by geographical proximity and broad institutional similarity (Table F.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 F.3 reports estimated cross-wage elasticities for the wage-quartile groups of countries, as discussed in Subsection 4.5.

Table F.1: RELATIVE DIFFERENCE OF UNCORRECTED AND CORRECTED INTENSIVE-MARGIN ESTIMATES

Relative difference in employment effect estimates		Wage change in HOM	CEE	DEV	OIN	WEU
		(1)	(2)	(3)	(4)	(5)
		Foreign affiliates in manufacturing				
HOM	*intensive*	-.081	.006	-.264	-.052	-.100
CEE	*intensive only*	-.081	-.043	-1.400	.083	-.148
DEV	*intensive only*	-.178	.102	-.065	.066	.075
OIN	*intensive only*	.013	-.010	.017	-.009	-.010
WEU	*intensive only*	-.121	.0007	.011	-.008	.117
		Foreign affiliates in any sector				
HOM	*intensive*	-.048	.015	.011	-.011	-.071
CEE	*intensive only*	-.095	-.060	-3.087	-.099	-.169
DEV	*intensive only*	-.019	-3.269	-.030	-4.244	-1.251
OIN	*intensive only*	-.009	.013	-4.352	-.017	-.043
WEU	*intensive only*	-.040	-.037	-1.268	-.013	-.039

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages), manufacturing MNEs and their majority-owned foreign affiliates in manufacturing (upper panel) or in any sector (lower panel).
Notes: The relative difference between elasticities at the intensive margin from uncorrected ISUR estimation and from parametric selectivity-corrected ISUR estimation (Assumption 1, Table 8) is the difference between the two elasticity estimates divided by selectivity-corrected estimates. For affiliates in manufacturing, there are 2,141 stacked MNE observations for uncorrected ISUR and 1,640 for selectivity-corrected ISUR estimation. For affiliates in any sector, there are 3,183 stacked MNE observations for uncorrected ISUR and 2,501 for selectivity-corrected ISUR estimation. Locations: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

Table F.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 F.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*	-.556***	.509***	.016	.029***	-.006
Qrtl. 4	*intensive only*	1.688***	-1.746***	.044	-.002	.008
	extensive only	.711***	-.786***	.048***	.017***	.002***
Qrtl. 3	*intensive only*	.245	.198	-.384	-.090**	.025
	extensive only	.711***	.214***	-.952***	.017***	.002***
Qrtl. 2	*intensive only*	1.175***	-.018	-.248**	-1.012***	.097***
	extensive only	.711***	.214***	.048***	-.983***	.002***
Qrtl. 1	*intensive only*	-1.693	.694	.485	.696**	-.188
	extensive only	.711***	.214***	.048***	.017***	-.998***

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages).
Notes: Elasticities at the extensive and intensive margins from 677 stacked MNE observations. Underlying labor demand estimates from parametric selectivity-corrected ISUR estimates (Assumption 1). Standard errors from 200 bootstraps: ** significance at five, *** one percent. Locations: HOM (Germany) and four foreign-country groups by manufacturing-wage quartiles, fourth quartile with top wages.

References

Amemiya, Takeshi, "Multivariate Regression and Simultaneous Equation Models when the Dependent Variables Are Truncated Normal," *Econometrica*, November 1974, *42* (6), 999–1012.

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.

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

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.

Brown, Randall S. and Laurits R. Christensen, "Estimates of Elasticities of Substitution in a Model of Partial Static Equilibrium: An Application to U.S. Agriculture, 1947-1974," in Ernst R. Berndt and Barry C. Field, eds., *Modeling and measuring natural resource substitution*, Cambridge, Massachusetts: MIT Press, 1981, chapter 10, pp. 209–29.

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.

Christensen, Laurits R., Dale W. Jorgenson, and Lawrence J. Lau, "Transcendental Logarithmic Production Frontiers," *Review of Economics and Statistics*, February 1973, *55* (1), 28–45.

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.

Diewert, W. E., "An Application of the Shephard Duality Theorem: A Generalized Leontief Production Function," *Journal of Political Economy*, May-June 1971, *79* (3), 481–507.

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.

Dixit, Avinash K., "Entry and Exit Decisions under Uncertainty," *Journal of Political Economy*, June 1989, *97* (3), 620–38.

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.

Feinberg, Susan E. and Michael P. Keane, "Accounting for the Growth of MNC-Based Trade Using a Structural Model of US MNCs," *American Economic Review*, December 2006, *96* (5), 1515–58.

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

Hall, Robert E., "The Specification of Technology with Several Kinds of Output," *Journal of Political Economy*, July-August 1973, *81* (4), 878–92.

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 John Ries, "Offshore Production and Skill Upgrading by Japanese Manufacturing Firms," *Journal of International Economics*, October 2002, *58* (1), 81–105.

_ **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.

Helpman, Elhanan, Marc Melitz, and Yona Rubinstein, "Estimating Trade Flows: Trading Partners and Trading Volumes," *Quarterly Journal of Economics*, May 2008, *123* (2), 441–87.

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*, 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.*

Meyerhoefer, Chad D., Christine K. Ranney, and David E. Sahn, "Consistent Estimation of Censored Demand Systems Using Panel Data," *American Journal of Agricultural Economics*, August 2005, *87* (3), 660–72.

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.

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

von Haefen, Roger H., Daniel J. Phaneuf, and George R. Parsons, "Estimation and Welfare Analysis with Large Demand Systems," *Journal of Business and Economic Statistics*, April 2004, *22* (2), 194–205.

Wales, Terence J. and A. D. Woodland, "Estimation of Consumer Demand Systems with Binding Non-Negativity Constraints," *Journal of Econometrics*, April 1983, *21* (3), 263–85.

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

Yen, Steven T. and Biing-Hwan Lin, "A Sample Selection Approach to Censored Demand Systems," *American Journal of Agricultural Economics*, August 2006, *88* (3), 7429.

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.