MARGINS OF MULTINATIONAL LABOR SUBSTITUTION

MARC-ANDREAS MUENDLER
SASCHA O. BECKER

CESIFO WORKING PAPER NO. 1713
CATEGORY 7: TRADE POLICY
MAY 2006

CESifo *Working Paper No. 1713*

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Abstract

Multinational labor demand responds to wage differentials at the extensive margin, when a multinational enterprise (MNE) expands into foreign locations, and at the intensive margin, when an MNE operates existing affiliates across locations. We derive conditions for parametric and nonparametric identification of an MNE model to infer elasticities of labor substitution at both margins, controlling for location selectivity. Prior studies have rarely found foreign wages or operations to affect employment. Our strategy detects salient adjustments at the extensive margin for German MNEs. With every percentage increase in German wages, German MNEs allocate 2,000 manufacturing jobs to Eastern Europe at the extensive margin and 4,000 jobs overall.

JEL Code: F21, F23, C14, C24, J23.

Keywords: multinational enterprise, location choice, sample selectivity, labor demand, translog cost function, nonparametric estimation.

Marc-Andreas Muendler
Department of Economics
University of California, San Diego
9500 Gilman Drive
La Jolla, CA 92093-0508
USA
muendler@ucsd.edu

Sascha Becker
Center for Economic Studies at the
University of Munich
Schackstr. 4
80539 Munich
Germany
sbecker@lmu.de

April 15, 2006

We thank Gordon Hanson, Xiaohong Chen, Peter Egger, Sebastian Kessing and Hal White as well as participants at various seminars and conferences for insightful suggestions. We thank Steve Redding for sharing code to compute market access statistics. Jennifer Poole, Robert Jäckle, Nadine Gröpl, and Daniel Klein provided excellent research assistance. Simone Hofer from UBS kindly shared the bank's international wage data. We gratefully acknowledge financial support from the VolkswagenStiftung under its grant initiative *Global Structures and Their Governance* and administrative and financial support from the Ifo Institute. Becker also gratefully acknowledges financial support from the Fritz-Thyssen-Stiftung.

1 Introduction

Multinational enterprises (MNEs) are important mediators of world trade. Surprisingly, however, the operation of MNEs has rarely been found to affect factor demands across locations (e.g. Slaughter (2000) for U.S., Konings (2004) for European MNEs). We quantify the effect of permanent wage differentials on MNE employment at two critical margins. An MNE's labor demand responds to international wage differentials at the *extensive margin*, when the MNE expands into a foreign market, and at the *intensive margin*, when the MNE operates existing affiliates and chooses employment. Our paper thus offers an integration of two strands of the empirical literature—one on MNEs' location choices (Devereux and Griffith 1998, Head and Mayer 2004) and one on MNE operations across existing locations (Slaughter 2000, Head and Ries 2002, Hanson, Mataloni, and Slaughter 2005)—into a unified estimation framework.

The MNE's two-stage decision, to first expand (extensive margin) and then operate (intensive margin), has a well-defined econometric counterpart in sample selection. Aside from the economic interpretation of the extensive margin, labor demand or cost function estimates at the intensive margin are subject to selectivity bias unless corrected. Using comprehensive data on German manufacturing MNEs and their majority-owned foreign manufacturing affiliates, we find that an MNE's propensity to select a foreign location is a salient predictor of its labor demand across locations and that permanent wage differentials have a strong impact on multinational labor substitution both at the extensive and the intensive margin.

A methodological contribution of our paper is to extend the univariate sample selection case to one of multiple selections. We derive conditions under which the common Heckman (1979) selection correction can be applied location by location to correct outcome estimation, in our case a seemingly unrelated equation system of the MNE's cost function. We also prove identification of a nonparametric selection model, which extends single-equation models (such as those in Das, Newey, and Vella (2003)) to the multivariate case. The nonparametric estimator is simple to implement in a two-stage approach and is applicable to the estimation of multivariate demand systems in general (for a recent parametric approach to multivariate demand see e.g. Yen (2005)).

To quantify the extensive margin, we base our parametric and non-parametric estimators of location selection on MNE-wide profit maximization. Existing firm-level studies on the expansion of MNEs do not find low wages or low per-capita incomes to be significant predictors of location choice (e.g. Devereux and Griffith (1998) for U.S., Head and Mayer (2004) for Japanese, Buch, Kleinert, Lipponer, and Toubal (2005) for German MNEs).[1] Multinomial logit estimation turns wages

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

into significant predictors of location choice in Disdier and Mayer (2004) for French MNEs, and in Becker, Ekholm, Jäckle, and Muendler (2005) for Swedish MNEs and the same German MNEs as in this paper. But multinomial logit estimation rests on the assumption that independent agents within the MNE decide on distinct investment projects; that is incompatible with MNE-wide profit maximization. Devereux and Griffith (1998) estimate multinomial logit choice and, to be consistent with MNE-wide optimization, restrict their sample to MNEs who invest in only one location abroad; they do not find wages to be significant predictors of U.S. MNEs' location choices. In contrast, when we condition on an MNEs' past presence and its interaction with wages, we find wage variables to be statistically significant predictors of location choices in probit and in non-parametric selection regressions. When weighted with the impact of location selection on employment, wage differentials across locations are substantial predictors of labor substitution within MNEs at the extensive margin.

At the intensive margin, 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.[2] Despite this apparent importance of MNEs for international transactions, Slaughter (2000) reports that, in a sample of U.S. MNEs, operations in low-wage locations have no detectable impact on MNE employment in the home market. In contrast, Feenstra and Hanson (1999) attributed about a third of U.S. relative wage changes to outsourcing (within MNEs or across firms). Similar to Slaughter (2000), Konings (2004) and Barba Navaretti and Castellani (2004) find no evidence for the hypothesis that operations of European MNEs in low-wage locations have an impact on home-market labor demand. Braconier and Ekholm (2000) and Marin (forthcoming) estimate wage elasticities of labor demand and intermediate imports from Central and Eastern Europe for Western European MNEs, and report no significant effect of foreign relative wages. Brainard and Riker (2001), however, do find that foreign affiliate employment substitutes modestly for U.S. parent employment but less so than for employment across foreign locations.[3] Hanson, Mataloni, and Slaughter (2005) shift focus from factor demands to intermediate input uses and, as an exception to most prior firm-level evidence, report that affiliates of U.S. MNEs process significantly more intra-firm imports the lower are low-skilled wages. The result challenges the view that relative abundance in low-skilled labor fails to attract MNEs. We revisit their result in the context of multinational labor substitution and extend the estimation framework to incorporate location choice. When controlling for the propensity to select a foreign location, wages are statistically significant and economically salient predictors of MNEs' labor demands at the intensive margin.

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activity).

[2] UNCTAD press release TAD/INF/PR/47 (12/08/02).

[3] At the aggregate level, Brainard (1997) does not find relative abundance of low-skilled labor to explain MNE sales patterns across locations.

Our findings point to large sunk entry and exit costs so that MNE expansions (or withdrawals) are infrequent but, when undertaken, they have a sizeable impact on labor demand. We find cross-wage elasticities at the extensive margin to be strictly positive. So, home and foreign employment are substitutes within MNEs not only at the intensive but also at the extensive margin. Elasticities at the extensive margin are about half the size of elasticities at the intensive margin in locations close to home. For overseas developing country wages, however, elasticities are significantly different from zero only at the extensive margin. Bootstraps reject equality between the intensive and the total elasticity of substitution for most locations, corroborating the importance of the extensive margin. Elasticity point estimates at both margins are robust across different samples and wage data, specifications, and parametric and nonparametric estimation techniques.

We evaluate the counterfactual question 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, 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 levels of wages, an increase in CEE wages of 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.[4] Of course, elasticities of substitution are not constant at all levels of wages 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 five more sections. Section 2 elaborates a model of the expansion and operation of MNEs, and Section 3 derives identification conditions for its estimation under location selectivity. Section 4 presents the data and discusses descriptive statistics on location choice. Estimation results on multinational labor substitution are presented in Section 5, and interpreted in counterfactual evaluations. Section 6 concludes.

2 Multinational Expansion and Operation

Let observed employment y_j^ℓ of MNE j at time t in location ℓ obey

$$y_{jt}^\ell = \mathbf{x}_{jt}^\ell \beta^\ell + \epsilon_{jt}^\ell$$

[4]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).

if MNE j is present at ℓ. Else, $y_{jt}^{\ell} = 0$. In the translog case, the vector $\mathbf{x}_{jt}^{\ell}$ of employment predictors includes additively separable transformations of outputs, inputs and factor prices (we discuss regressor construction below), including the prevailing wage differentials between locations at time t. ϵ_{jt}^{ℓ} is a disturbance term. So, the conditional expectation of MNE j's observed employment in location ℓ is

$$\bar{y}_{jt}^{\ell} \equiv \mathbb{E}\left[y_{jt}^{\ell} \,|\, \mathbf{x}_{jt}^{\ell}, \mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}\right] = \mathbf{x}_{jt}^{\ell}\beta^{\ell} + \mathbb{E}\left[\epsilon_{jt}^{\ell} \,|\, \mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}\right], \tag{1}$$

where the vector $\mathbf{d}_{jt}$ of presence indicators d_{jt}^{k} reflects MNE j's observed pattern of locations $k = 1, \ldots, L$ at time t ($d_{jt}^{k} = 1$ if firm j is present in location k and $d_{jt}^{k} = 0$ otherwise) and contains $d_{jt}^{\ell} = 1$. The information set $\mathbf{z}_{j,t-\tau}$ at moment $t - \tau$ affects labor demand through the resulting choice of presence in location ℓ.

We define the *extensive* margin of labor demand to be the expected labor demand $\bar{y}_{\ell}^{\text{ext}}$ in location ℓ, predicted by a firm j's current choices of presence around the world and its past information set $\mathbf{z}_{j,t-\tau}$,

$$\bar{y}_{jt}^{\text{ext},\ell} \equiv \mathbb{E}\left[\epsilon_{jt}^{\ell} \,|\, d_{jt}^{1}, \ldots, d_{jt}^{\ell} = 1, \ldots, d_{jt}^{L}; \mathbf{z}_{j,t-\tau}\right], \tag{2}$$

where the optimal binary choices $(d_{jt}^{1}, \ldots, d_{jt}^{\ell}, \ldots, d_{jt}^{L})$ are functions of MNE j's information set at the moment of location choice $t - \tau$, and τ is the time it takes an MNE to implement location choices (two to four years, say). The information set $\mathbf{z}_{j,t-\tau}$ at moment $t-\tau$ predicts presence in location k with $d_{jt}^{k} = \mathbf{1}(H(\mathbf{z}_{j,t-\tau}) + \eta_{j,t-\tau}^{k} > 0)$, where $H(\cdot)$ is an unknown function and $\eta_{j,t-\tau}^{k}$ is a disturbance to the MNE's presence. Most important, $\mathbf{z}_{j,t-\tau}$ includes the then prevailing wage differentials between locations.

Labor demand at the *intensive* margin is accordingly defined as

$$\bar{y}_{jt}^{\text{int},\ell} \equiv \bar{y}_{jt}^{\ell} - \bar{y}_{jt}^{\text{ext},\ell} = \mathbf{x}_{jt}^{\ell}\beta^{\ell}. \tag{3}$$

The labor demand effect at the extensive margin $\bar{y}_{jt}^{\text{ext},\ell} = \mathbb{E}[\epsilon_{jt}^{\ell} \,|\, \mathbf{d}_{jt}]$ is an additive component of conditional labor demand $\mathbb{E}[y_{jt}^{\ell} \,|\, \mathbf{x}_{jt}^{\ell}, \mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}]$. Economically, an MNE's mere presence at a location typically raises the labor demand prediction for that location.[5] Statistically, the extensive margin needs to be included in the regression to correct for selectivity.

MNE j produces a vector of location-specific outputs $\mathbf{q}_{jt} = (q_{jt}^{1}, \ldots, q_{jt}^{L})'$ at L locations. We consider MNEs to be price takers in input market, whereas they may have market power in output markets. (We estimate a cost function, so any pricing behavior in the sales market is consistent with our approach.) On the input side, we focus on employment. We view MNEs as wage takers in the local markets, competing with labor demand from non-tradeable goods sectors and incumbent

[5] To be precise, this is true if high home wages raise the probability of presence at a foreign location ℓ and the presence likelihood is positively correlated with labor demand at that foreign location ℓ. Both conditions are satisfied in our MNE sample.

firms. Similarly, we consider demand for capital goods and intermediate inputs from non-MNEs as sufficiently large so that the remaining demand of MNEs for those goods has a negligible price impact.

Final goods prices are world-market prices that differentiated products *from* locations $\ell = 1, \ldots, L$ can fetch, given product characteristics. Final goods are produced with labor and capital. After controlling for location choice in the formation of the MNE, we consider installed capital $\mathbf{k}_{jt} = (k^1_{jt}, \ldots, k^L_{jt})'$ to be a quasi-fixed factor in an MNE's short-run cost function C_{jt} (but put to use at locations $k = 1, \ldots, L$ to different degrees). We consider labor at locations $k = 1, \ldots, L$ to be immobile across national borders and its factor prices $\mathbf{w}_t = (w^1_t, \ldots, w^L_t)'$ as specific to L locations.

2.1 Location choice

Define γ^ℓ_N as the fixed FDI entry costs at location ℓ and γ^ℓ_X as the fixed FDI exit costs from location ℓ.[6] Then, fixed costs of changing presence at location ℓ in t, anticipated at $t - \tau$, become

$$G^\ell(d^\ell_{jt}, d^\ell_{j,t-\tau}) = \gamma^\ell_N \, d^\ell_{jt}(1-d^\ell_{j,t-\tau}) + \gamma^\ell_X \,(1-d^\ell_{jt})d^\ell_{j,t-\tau},$$

where d^ℓ_{jt} is the indicator for MNE j's current FDI presence at location ℓ, and $d^\ell_{j,t-\tau}$ for its past presence. We restrict the long-term fixed cost components γ^ℓ_N and γ^ℓ_X to be time invariant in our four-year MNE panel data (but control for time-varying country and MNE characteristics in selection estimation). The decision-relevant fixed cost difference $F^\ell_{j,t-\tau} \equiv G^\ell(1, d^\ell_{j,t-\tau}) - G^\ell(0, d^\ell_{j,t-\tau})$ between presence at location ℓ and absence from ℓ at time t is

$$F^\ell_{j,t-\tau} = \gamma^\ell_N - (\gamma^\ell_X + \gamma^\ell_N)\, d^\ell_{j,t-\tau}, \tag{4}$$

where $(\gamma^\ell_X + \gamma^\ell_N)$ is sometimes called the *hysteresis band* and reflects the sunk cost effect that induces firms to continue operations at location ℓ (Dixit 1989).[7]

To select locations (τ years prior to production and sales), MNE j maximizes expected profits $\mathbb{E}_{j,t-\tau}[\,\mathbf{p}(\mathbf{q}_{i\neq j,t}, \mathbf{q}_{jt})' \cdot \mathbf{q}_{jt} - C_{jt}(\mathbf{q}_{jt}; \mathbf{k}_{jt}, \mathbf{w})]$. This implies that MNE j's rule for FDI presence at location ℓ can be written as

$$\begin{aligned}
d^\ell_{jt} &= \mathbf{1}\Big(\mathbb{E}_{j,t-\tau}[p^\ell q^{\ell,*}_{jt}] + \mathbb{E}_{j,t-\tau}[C_{jt}(q^\ell_{jt}{=}0; \cdot) - C_{jt}(q^{\ell,*}_{jt}; \cdot)] - F^\ell_{j,t-\tau} + \eta^\ell_{j,t-\tau} > 0\Big) \\
&= \mathbf{1}\left(h(\mathbf{z}^0_{j,t-\tau}) - \gamma^\ell_N + (\gamma^\ell_X + \gamma^\ell_N)\, d^\ell_{j,t-\tau} + \eta^\ell_{j,t-\tau} > 0\right) \\
&= \mathbf{1}\left(H(\mathbf{z}_{j,t-\tau}) + \eta^\ell_{j,t-\tau} > 0\right)
\end{aligned} \tag{5}$$

[6] For simplicity, the fixed costs of reentry into a given location after a period of absence are assumed to be equal to the costs at first entry γ^ℓ_N.

[7] Probit estimation with firm-fixed effects is known for problematic performance in panel data with a short time horizon (Heckman 1981). We therefore do not attempt to estimate MNE-specific sunk costs of presence $F^\ell_{j,t-\tau}$ at location ℓ. We distinguish between entry and exit sunk cost components to account for MNE-specific differences in $F^\ell_{j,t-\tau}$, similar to Roberts and Tybout's (1997) model of sunk costs in exporting status.

(see Appendix A for a derivation). The unknown function $h(\mathbf{z}^0_{j,t-\tau})$ captures both expected revenues from producing the profit-maximizing quantity $q^{\ell,*}_{jt}$ at location ℓ and expected cost savings from producing at ℓ (see first line). Sunk costs of presence at location ℓ have an observable component $F^{\ell}_{j,t-\tau}$ by (4) and a disturbance $\eta^{\ell}_{j,t-\tau}$. The disturbance $\eta^{\ell}_{j,t-\tau}$ is known to the MNE but not to the researcher. To simplify notation, we write $H(\mathbf{z}_{j,t-\tau}) \equiv h(\mathbf{z}^0_{j,t-\tau}) - \gamma^{\ell}_N + (\gamma^{\ell}_X + \gamma^{\ell}_N)\, d^{\ell}_{j,t-\tau}$ and include past presence in any location in the information set $\mathbf{z}_{j,t-\tau}$.

Equation (5) is the selection equation: the empirical rule of presence in locations $\ell = 1, \ldots, L$. We estimate the rule both parametrically (with a probit regression and $H(\mathbf{z}_{j,t-\tau}) = \mathbf{z}_{j,t-\tau}\gamma^{\ell}$) and nonparametrically.

2.2 Multiproduct cost function

To obtain theoretically well-defined estimates of elasticities of labor substitution across locations, we opt for a flexible parametric specification of the MNE's multiproduct cost function. We first augment the cost function with parametric corrections for location selectivity. We then proceed to a model with a parametric cost function part and a nonparametric correction for selectivity.

We use a short-run multiproduct translog cost function to estimate labor demand, and extend it to control for location selectivity.[8] A short-run cost function, given MNE j's location choice, treats MNE j's vector of capital stocks $\mathbf{k}_{jt}$ as quasi-fixed factors. We prefer a short-run over a long-run cost function because we already control for the installation of foreign affiliates through location selectivity (5) and because the inclusion of capital stock variables captures otherwise unobservable (firm-specific) user costs of capital across locations.

Applying Shepard's (1953) lemma to the short-run multiproduct translog cost function yields location-specific wage bill shares $s^{\ell}_{jt} \equiv w^{\ell}_t y^{\ell}_{jt}/C_{jt}$ (the wage bill at location ℓ in the MNE's total wage bill) as functions of $(\mathbf{q}_{jt}; \mathbf{k}_{jt}, \mathbf{w})$. We multiply the wage bill shares s^{ℓ}_{jt} with observation-specific scalars C_{jt}/w^{ℓ}_t to arrive at our outcome equation (labor demand at ℓ)

$$y^{\ell}_{jt} = \mathbf{x}^{\ell}_{jt}\,\beta^{\ell} + \epsilon^{\ell}_{jt} \tag{6}$$

with

$$\mathbf{x}^{\ell}_{jt}\beta^{\ell} = \alpha_{\ell}\frac{C_{jt}}{w^{\ell}_t} + \sum_{m=1}^{L}\Big(\mu_{\ell m}\ln\Big[(q^m_{jt})^{C_{jt}/w^{\ell}_t}\Big] + \kappa_{\ell m}\ln\Big[(k^m_{jt})^{C_{jt}/w^{\ell}_t}\Big] + \delta_{\ell m}\ln\Big[(w^m_t)^{C_{jt}/w^{\ell}_t}\Big]\Big)$$

[8] We follow Brown and Christensen's (1981, eq. 10.21) short-run version of Christensen, Jorgenson, and Lau (1973) and extend the framework to multiple products. A main alternative would be Hall's (1973) generalization of Diewert's (1971) Leontief cost function to the multiproduct case. We favor the translog cost function because its dimensionality requirements are considerably leaner and permit higher-order approximations to the nonparametric correction for selectivity. Kohli (1978) took the translog specification to the empirical trade literature.

(see Appendix B), where ϵ^{ℓ}_{jt} is a disturbance.

Compared to translog regression equations in wage bill shares s^{ℓ}_{jt}, the transformation with observation-specific scalars C_{jt}/w^{ℓ}_{t} to an equivalent regression of y^{ℓ}_{jt} on $\mathbf{x}^{\ell}_{jt}$ has three important advantages. First, 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. Second, wages are regressors only and do not enter the dependent variable. Third, labor demand is not bounded above so that, conditional on $\mathbf{x}^{\ell}_{jt}$, the labor demand disturbance satisfies the assumption of a one-sided truncation for (parametric and nonparametric) selectivity correction.

Stacking locations with zero output and factor use. Most MNEs produce in some but not in all locations. For cases of zero output or input, however, equation (6) is not well defined. Especially zero turnover and zero capital stocks require attention because they are MNE-specific, but absence from a location also suggests dropping wage regressors when no employment occurs.

One possible treatment is estimation of separate equation systems for every single presence pattern in the data. The resulting estimators are hard to interpret, however, and plagued by dimensionality: potential presence in up to $L-1$ locations outside the home location implies that there are up to $2^{L-1}-1$ regional presence patterns for an MNE.[9] In the German sample in 2000, for instance, only 57 out of 1,770 MNEs are omnipresent in all four world locations while every single one of the 15 possible regional presence pattern occurs. So, there would be 15 sets of estimates.

We choose to stack observations of all MNEs in the sample. Stacking observations 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. Stacking is permissible under three conditions: (i) all MNEs face identical sunk cost $F^{\ell}_{j,t-\tau}$ for presence at location ℓ conditional on their prior presence and information set (so that presence is not correlated with inputs); (ii) MNEs face an identical short-run cost function $C_{jt}(\cdot) = C(\cdot)$ in all locations of presence, conditional on their characteristics (so that one common parameter vector is justified); and (iii) the disturbances ϵ^{ℓ}_{jt} are uncorrelated across observations.

We set all missing location variables for an absent MNE j to zero—that is log employment, turnover, capital stock and wages are zero at location m from where MNE j is absent. This is equivalent to interacting the translog cost function coefficients with presence indicators: $\mu_{\ell m} = 0$ when no output is produced at location m, and $\kappa_{\ell m} = \delta_{\ell m} = 0$ when MNE j employs no factors at location m. Stacking can induce correlations between the transformed regressors and the error ϵ^{ℓ}_{jt} in (6).

[9] MNEs are present in their home location by sample definition, so only 2^{L-1} patterns are observable in principle. Firms that only operate domestically without any foreign affiliate are not MNEs by definition so that the single presence pattern with the only presence at the home location must be subtracted.

To remove this source of potential bias, we include the set of absence indicators $(\mathbf{1}-\mathbf{d}_{jt})$ (with nuisance parameters β^{ℓ}_{d}) among the regressors in the outcome equation: $y^{\ell}_{jt} = \mathbf{x}^{\ell}_{jt}\beta^{\ell} = \mathbf{x}^{0\ell}_{jt}\,\beta^{\ell} + (\mathbf{1}-\mathbf{d}_{jt})\,\beta^{\ell}_{d}$. The set of absence indicators $(\mathbf{1}-\mathbf{d}_{jt})$ also offsets the zero output prediction at the sample mean.

3 Estimation under Location Selectivity

The selection equation (5) for location ℓ is

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

and, conditional on MNE j's selection of location ℓ, expectations of the outcome (6) are

$$\mathbb{E}\left[y^{\ell}_{jt}\,|\,\mathbf{x}^{\ell}_{jt}, \mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}\right] = \mathbf{x}^{\ell}_{jt}\beta^{\ell} + \mathbb{E}\left[\epsilon^{\ell}_{jt}\,|\,d^{1}_{jt}, \ldots, d^{\ell}_{jt} = 1, \ldots, d^{L}_{jt}; \mathbf{z}_{j,t-\tau}\right],$$

where disturbances ϵ^{ℓ}_{jt} and $\eta^{\ell}_{j,t-\tau}$ are uncorrelated across observations (of MNEs i and j, and between periods t and $t+1$). The timing of $\eta^{\ell}_{j,t-\tau}$ is not important and the $\eta^{\ell}_{j,t-\tau}$ realization could be simultaneous with ϵ^{ℓ}_{jt}. Natural exclusion restrictions on covariates that do not enter the cost function identify location selection.

In this section, we discuss cross-regional distributional assumptions on $(\epsilon^{\ell}_{jt}, \eta^{\ell}_{j,t-\tau})$ and permissible estimation techniques under those conditions. For a parametric cost function specification (with well-defined elasticities of substitution), a parametric approach to selectivity appears natural to start with. We present sets of necessary and sufficient distributional assumptions for univariate Heckman (1979) corrections location by location, to which we refer as *parametric* selectivity correction. Empirical evidence on the necessary assumptions is favorable in our sample. For multivariate selectivity, an extension of the 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.

To be free of distributional restrictions, we extend the parametric approach to a nonparametric multivariate selection model (similar to one in Das, Newey, and Vella (2003)) and account for cross-location correlations between labor demand choices at the extensive and intensive margins. We derive identification from common sufficient assumptions. The nonparametric procedure allows for unknown disturbance distributions, and for unknown functional forms of $\mathbb{E}\left[\epsilon^{\ell}_{jt}\,|\,\mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}\right]$ and $\mathbf{1}(H(\mathbf{z}_{j,t-\tau}) + \eta^{k}_{j,t-\tau} > 0)$.

3.1 Parametric selectivity correction

Consider Heckman (1979) selectivity corrections location by location. There are two alternative sets of assumptions that allow for such a parametric correction, whereby labor demand (6) in ℓ only requires correction for selectivity (5) into ℓ but not into any other locations $k \neq \ell$. We are interested in $\mathbb{E}\left[y^{\ell}_{jt} \,|\, \mathbf{x}^{\ell}_{jt}, \mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}\right]$ and $H(\mathbf{z}_{j,t-\tau}) = \mathbf{z}_{j,t-\tau}\gamma^{\ell} - \gamma^{\ell}_N + (\gamma^{\ell}_X + \gamma^{\ell}_N)d^{\ell}_{j,t-\tau}$.

Assumption 1 *The disturbances $(\epsilon^k_{jt}, \eta^{\ell}_{j,t-\tau})$ are multivariate normally distributed and independent of $\mathbf{x}^m_{jt}$ and $\mathbf{z}_{j,t-\tau}$ for all k, ℓ, m (and $\mathbb{V}ar(\eta^{\ell}_{j,t-\tau}) = 1$). In addition, either*

(a) the part of the selection shock that correlates with labor demand shocks is an MNE-specific disturbance and does not vary by location so that, conditional on the MNE-specific shocks, ϵ^k_{jt} and ϵ^{ℓ}_{jt} as well as $\eta^k_{j,t-\tau}$ and $\eta^{\ell}_{j,t-\tau}$ are independent for $k \neq \ell$, or

(b) the labor-demand related part of the selection shock varies by location but is independent of labor demand shocks in other locations (ϵ^k_{jt} and $\eta^{\ell}_{j,t-\tau}$ are independent for $k \neq \ell$),

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

Especially case (a), where the part of the selection shock $\eta^{\ell}_{j,t-\tau}$ that correlates with labor demand shocks ϵ^k_{jt} is an MNE-specific disturbance and does not vary by location, is plausible in economic terms. Suppose selection disturbances include both host country-specific parts such as, for example, surprising changes to profit repatriation policies and include MNE-specific parts such as shocks to its 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-relevant host-country shocks so that cost function disturbances are unrelated to the entry-relevant policy shocks. In case (a), all relevant information for labor demand at any location ℓ is fully contained in the single indicator d^{ℓ}_{jt} (which is as informative about $\eta^{\ell}_{j,t-\tau}$ as any other location indicator). Case (b) is more restrictive and implies that neither MNE-specific nor host-country specific shocks to presence at location ℓ have a bearing on labor demand at other locations $k \neq \ell$.

Note that cross-location correlations of labor demand shocks are not necessarily evidence against Assumption 1. As the proof to Proposition 1 will show, case (a) of MNE-specific selection shocks induces a correlation between labor demand shocks across locations: ϵ^k_{jt} and $\eta^{\ell}_{j,t-\tau}$ correlate across locations $k \neq \ell$ but in the same way as ϵ^{ℓ}_{jt} and $\eta^{\ell}_{j,t-\tau}$.

Proposition 1 *Independent parametric selection correction for L locations identify $\mathbf{x}^{\ell}_{jt}\beta^{\ell}$ and $\mathbb{C}ov(\epsilon^{\ell}_{jt}, \eta^{\ell}_{j,t-\tau})$ if and only if Assumption 1 holds.*

Proof. Because any normally distributed variable can be linearly decomposed into a sum of independent standard normal variables, consider without loss of generality

$$\eta^{\ell}_{j,t-\tau} = \sqrt{1-\omega}\, e^{\ell}_{jt} + \sqrt{\omega}\, \sum_{k\leq\ell} \frac{\pi^{k\ell}_{\eta}}{\sqrt{\sum_{k\leq\ell}(\pi^{k\ell}_{\eta})^2}}\, u^k_{jt}, \tag{7}$$

$$\epsilon^{\ell}_{jt} = \sum_k \lambda^{k\ell}\, e^k_{jt} + \sum_{k\leq\ell} \pi^{k\ell}_{\epsilon}\, \upsilon^k_{jt} \tag{8}$$

for independent standard normal variables $e^k_{jt}, u^k_{jt}, v^k_{jt}$ $(k = 1, \ldots, L)$, where $\omega \in [0, 1]$ is a weight to satisfy $(\sigma^\ell_\eta)^2 = \sigma^{\ell\ell}_\eta = 1$, and $\pi^{k\ell}_\eta$, $\pi^{k\ell}_\epsilon$, $\lambda^{k\ell}$ are parameters. To prove sufficiency, let $\pi^{k\ell}_\eta = \pi^{k\ell}_\epsilon = 0$ for $k \neq \ell$.

First consider (a) MNE-specific selection shocks $\eta^\ell_{j,t-\tau}$ whose labor demand related part does not vary over locations. Concretely, set $e^k_{jt} = e_{jt}$ for all locations k, and denote $\lambda^{\cdot\ell} \equiv \sum_k \lambda^{k\ell}$. Then the variances and covariances of the selection shocks (7) are $\sigma^{\ell\ell}_\eta = 1$ and $\sigma^{k\ell}_\eta = 1-\omega$. The variances and covariances of the labor demand shocks (8) are $\sigma^{\ell\ell}_\epsilon = (\lambda^{\cdot\ell})^2 + (\pi^{\ell\ell}_\epsilon)^2$ and $\sigma^{k\ell}_\epsilon = (\lambda^{\cdot\ell})^2$. And the covariances between the selection shock in location k and the demand shock in location ℓ are $\sigma^{k\ell}_{\eta\epsilon} = \lambda^{\cdot\ell}$.

Second, consider (b) location-varying selection shocks $\eta^\ell_{j,t-\tau}$ that are independent of labor demand shocks in other locations. Concretely, set $\lambda^{k\ell} = 0$ for $k \neq \ell$, and denote $\lambda^{\cdot\ell} \equiv \lambda^{\ell\ell}$ for comparability. Then the selection shock variances and covariances are $\sigma^{\ell\ell}_\eta = 1$ and $\sigma^{k\ell}_\eta = 0$. The variances and covariances of the labor demand shocks are $\sigma^{\ell\ell}_\epsilon = (\lambda^{\cdot\ell})^2 + (\pi^{\ell\ell}_\epsilon)^2$ and $\sigma^{k\ell}_\epsilon = 0$. The covariances between the selection shock in location k and the demand shock in location ℓ are $\sigma^{\ell\ell}_{\eta\epsilon} = \sqrt{1-\omega}\,\lambda^{\cdot\ell}$ and $\sigma^{k\ell}_{\eta\epsilon} = 0$ for $k \neq \ell$.

In both cases, the marginal likelihood function becomes

$$g(y^\ell_{jt}|\mathbf{x}^\ell_{jt}, \mathbf{z}_{j,t-\tau}) = \frac{\phi\left((y^\ell_{jt} - \mathbf{x}^\ell_{jt}\beta^\ell)/\sigma^\ell_\epsilon\right)}{\sigma^\ell_\epsilon\,\Phi(\mathbf{z}_{j,t-\tau}\gamma^\ell)} \cdot \Phi\left(\frac{\rho^{\ell\ell}_{\eta\epsilon}(y^\ell_{jt} - \mathbf{x}^\ell_{jt}\beta^\ell) + \mathbf{z}_{j,t-\tau}\gamma^\ell}{\sigma^\ell_\epsilon\,(1-\rho^{\ell\ell}_{\eta\epsilon})^{1/2}}\right), \tag{9}$$

after concentrating out u^ℓ_{jt} and v^ℓ_{jt}, where $\sigma^\ell_\epsilon = \sqrt{\sigma^{\ell\ell}_\epsilon}$ and $\rho^{\ell\ell}_{\eta\epsilon} = \sigma^{\ell\ell}_{\eta\epsilon}/\sigma^\ell_\epsilon$, and $\phi(\cdot)$ and $\Phi(\cdot)$ are the standard normal density and distribution functions. This is precisely the likelihood function for independent Heckman (1979) correction location by location.

For necessity, observe that parameters $\pi^{k\ell}_\eta \neq 0$ or $\pi^{k\ell}_\epsilon \neq 0$ for any $k \neq \ell$ cause cross-equation correlations and do not permit concentrating out u^ℓ_{jt} and v^ℓ_{jt} to arrive at (9). Similarly, $\lambda^{k\ell} \neq 0$ for any $k \neq \ell$ precludes concentrating out e^ℓ_{jt} to arrive at (9). ■

Estimation. Extending the parametric two-stage procedure to L locations, we first estimate equations (5) with probit regressions by location. Second, we estimate outcome (6) at location ℓ by including the predicted selectivity hazard (inverse of the Mills ratio) $\hat{\Lambda}^\ell_{jt}$ from the first stage among the regressors (we also include absence indicators $(\mathbf{1}-\mathbf{d}_{jt})$ among the regressors to prevent stacking bias). The coefficient on the predicted selectivity hazard equals $\beta^\ell_\Lambda \equiv \rho^{\ell\ell}_{\epsilon\eta}\sigma^\ell_\epsilon$. We implement the second-stage estimation of (6) for $L-1$ locations (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 L (Barten 1969). Through constraints, we impose linear homogeneity in factor prices and symmetry of wage coefficients (see appendix B). We treat induced heteroskedasticity following Heckman (1979) (resulting in differing standard errors on symmetric coefficients). After estimation, we test whether either of the two possible sets of distributional assumptions are satisfied. We will find implications of set (b) violated but fail to find evidence against (a).

Tests. Implications of Assumption 1 are testable. In case (a) of MNE-specific selection shocks and for any $\omega < 1$, Assumption 1 implies that $\sigma_\eta^{k\ell}$ is the same for any pair of locations $k \neq \ell$. Note that we have no evidence on $\sigma_{\eta\epsilon}^{k\ell}$ for $k \neq \ell$ from location-by-location estimation. We obtain estimates of $\sigma_\eta^{k\ell}$ from multivariate probit estimation instead and use a χ^2-test for their equality.

Under the additional assumption that $\omega = 0$, there is a further test to query case (a), whether selection shocks are purely MNE-specific. Probit (maximum likelihood) estimation of selection in the Heckman procedure does not predict the disturbances η_{jt}. A testable implication of an MNE-specific selection shock, however, is that, if an MNE is neither present in all locations nor absent from all locations, the choices of presence and absence must be consistent with a location-independent MNE-specific selection shock for all locations. Concretely, an MNE observation contradicts the assumption of a location-independent selection shock if $\mathbf{z}_{j,t-\tau}\gamma^k - F_{j,t-\tau}^k > \mathbf{z}_{j,t-\tau}\gamma^\ell - F_{j,t-\tau}^\ell$ for locations k of absence and locations ℓ of presence because η_{jt} can be subtracted from both sides of the inequalities. This implication is testable for the predicted values, which are normally distributed conditional on $\mathbf{z}_{j,t-\tau}$ and $\mathbf{d}_{j,t-\tau}$ by normality of η_{jt}.

For (b) location-variant selection shocks, the set of assumptions implies that $\sigma_\epsilon^{k\ell} = 0$. So, a regression of ϵ_{jt}^ℓ on $\epsilon_{jt}^1, \ldots, \epsilon_{jt}^{\ell-1}, \epsilon_{jt}^{\ell+1}, \ldots, \epsilon_{jt}^L$ must have zero coefficients. We test this implication.

Both sets (a) and (b) of assumptions imply that ϵ_{jt}^k is independent of d_{jt}^k for all k because ϵ_{jt}^k and $\eta_{j,t-\tau}^\ell$ are independent. We include absence indicators ($\mathbf{1}$–$\mathbf{d}_{jt}$) among the regressors in the outcome equation, however, so this is not a useful implication in our context.

3.2 Nonparametric selectivity correction

In the nonparametric version of the multivariate binary choice model (5) and (6),

$$\begin{aligned} d_{jt}^\ell &= \mathbf{1}\left(H(\mathbf{z}_{j,t-\tau}) + \eta_{j,t-\tau}^\ell > 0\right), \qquad (\ell = 1, \ldots, L) \\ \mathbb{E}\left[y_{jt}^\ell \,|\, \mathbf{x}_{jt}^\ell, \mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}\right] &= \mathbf{x}_{jt}^\ell \beta^\ell + \mathbb{E}\left[\epsilon_{jt}^\ell \,\middle|\, d_{jt}^\ell = 1, \mathbf{d}_{jt}^{k\neq\ell}; \mathbf{z}_{j,t-\tau}\right], \end{aligned}$$

no distributional assumptions are placed on $\eta_{j,t-\tau}^\ell$ or ϵ_{jt} and $H(\cdot)$ is an unknown function.

We augment the nonparametric sample selection model in Das, Newey, and Vella (2003) to remain identified under multivariate binary selection (similar in spirit to a selection model with endogeneity in Das, Newey, and Vella (2003)). Suppose $\eta_{j,t-\tau}^k$

and ϵ^ℓ_{jt} are correlated. Suppose also that $\mathbf{z}_{j,t-\tau}$ and $\mathbf{x}^\ell_{jt}$ are correlated (e.g. wages in the past and present, as our data show). Because d^k_{jt} is a function of $\eta^k_{j,t-\tau}$, it correlates with ϵ^ℓ_{jt}; because d^k_{jt} is a function of $\mathbf{z}_{j,t-\tau}$, it correlates with $\mathbf{x}^\ell_{jt}$. So, if the labor demand equation does not condition on d^k_{jt}, the identifying restriction that $\mathbf{x}^\ell_{jt}$ and y^ℓ_{jt} are uncorrelated will be violated.

Define the *propensity score* (the expected probability of selection conditional on $\mathbf{z}_{j,t-\tau}$) as $p^\ell_{jt} \equiv \mathbb{E}[d^\ell_{jt} \,|\, \mathbf{z}_{j,t-\tau}] = 1-G(-H(\mathbf{z}_{j,t-\tau}))$, where $G(\cdot)$ is the cumulative distribution function of $\eta^\ell_{j,t-\tau}$. Then, assuming $G(\cdot)$ is one-to-one and changing variables with $u^\ell_{jt} = 1-G(\eta^\ell_{j,t-\tau})$, labor demand at the extensive margin becomes

$$\begin{aligned}
\mathbb{E}[\epsilon^\ell_{jt} \,|\, d^\ell_{jt} = 1, \mathbf{d}^{k\neq\ell}_{jt}, \mathbf{z}_{j,t-\tau}] &= \mathbb{E}[\epsilon^\ell_{jt} \,|\, \eta^\ell_{j,t-\tau} > -H(\cdot); \mathbf{d}^{k\neq\ell}_{jt}, \mathbf{z}_{j,t-\tau}] \\
&= \mathbb{E}[\epsilon^\ell_{jt} \,|\, u^\ell_{jt} < p^\ell_{jt}; \mathbf{d}^{k\neq\ell}_{jt}] \\
&= \int \int_0^{p^\ell_{jt}} \epsilon^\ell_{jt}\, f(\epsilon^\ell_{jt}, u^\ell_{jt} | \mathbf{d}^{k\neq\ell}_{jt})\, \mathrm{d}\epsilon^\ell_{jt} \mathrm{d}u^\ell_{jt} \,/ p^\ell_{jt} \\
&= m^\ell\left(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt}\right).
\end{aligned}$$

So, the conditional labor demand disturbance for location ℓ depends only on the propensity score for that location and the pattern of presence elsewhere. Observed labor demand then satisfies

$$\mathbb{E}\left[y^\ell_{jt} \,|\, \mathbf{x}^\ell_{jt}, \mathbf{d}_{jt}, \mathbf{z}_{j,t-\tau}\right] = \mathbf{x}^\ell_{jt}\beta^\ell + m^\ell\left(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt}\right).$$

To establish identification, consider deviations from the truth $\Delta\xi^\ell(\mathbf{x}^\ell_{jt}) \equiv \mathbf{x}^\ell_{jt}(\hat{\beta}^\ell - \beta^\ell)$ and $\Delta m^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt}) \equiv \hat{m}^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt}) - m^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt})$, where hats denote estimates of the true (not hatted) functions. Assumption 2 states sufficient conditions for identification.

Assumption 2

(i) $\mathbb{E}[\epsilon^\ell_{jt} \,|\, d^\ell_{jt} = 1, \mathbf{d}^{k\neq\ell}_{jt}, \mathbf{z}_{j,t-\tau}] = m^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt})$,

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

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

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

Part (i) requires, as in the parametric case, that the conditional expectation of the labor demand disturbance at location ℓ is only a function of the propensity score of presence at ℓ and observed presence elsewhere. So, in the regression of observed labor demand y^ℓ_{jt} on $\mathbf{x}^\ell_{jt}\beta^\ell$ and $m^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt})$, $\mathbf{x}^\ell_{jt}\beta^\ell$ is a separate additive component. This specification extends nonparametric selectivity correction in Das, Newey, and Vella (2003) to the multivariate case.

Part (ii) is the same identification condition as in Das, Newey, and Vella (2003) and implies that p^ℓ_{jt} (which enters $m^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt})$) 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 $\Delta\xi^\ell(\mathbf{x}^\ell_{jt}) = m^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt})$ and $\Delta m^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt}) = -m^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt})$ indeterminate—a violation of (ii). In our context, the exclusion restriction arises naturally because the MNE chooses $\mathbf{x}^\ell_{jt}$ in response to information after $t-\tau$, whereas the decision of presence is based on $\mathbf{z}_{j,t-\tau}$. In addition, parent-firm characteristics and competitor-level host-country characteristics are predictors of presence but not related to the labor-cost 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.

Assumption 2 allows us to relax the earlier identifying assumption that $(\epsilon^k_{jt}, \eta^\ell_{j,t-\tau})$ is independent of $\mathbf{x}^m_{jt}$ and $\mathbf{z}_{j,t-\tau}$ for all k, ℓ, m. Assumption 2 only requires that, conditional on the propensity score p^ℓ_{jt}, ϵ^ℓ_{jt} is uncorrelated with all functions of $\mathbf{x}^\ell_{jt}$ and $\mathbf{z}_{j,t-\tau}$. Moreover, the nonparametric estimator $\mathbf{x}^m_{jt}$ allows for conditional heteroskedasticity of unknown form (and thus presents a nonparametric alternative to Chen and Khan's (2003) three-step estimator). Also note that we need no assumption on the cross-equation correlation of $\eta^\ell_{j,t-\tau}$ if we include $\mathbf{d}^{k\neq\ell}_{jt}$. This makes nonparametric analysis a powerful tool for multivariate binary selection estimation.

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

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

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

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

Note that lacking identification of additive constants is not a problem in our context. The transformed cost function regressors $\mathbf{x}^\ell_{jt}\beta^\ell$ in equation (6) do not include a constant term. To assess the labor demand effect of permanent wage differentials at the extensive margin, we will evaluate $\nabla_{p^\ell_{jt}} m^\ell(p^\ell_{jt}, \mathbf{d}^{k\neq\ell}_{jt}) \cdot \nabla_{\mathbf{z}_{j,t-\tau}} p^\ell_{jt}$ (a scalar), for which the constant does not matter.

Conversely, if we want to include the propensity scores $\mathbf{p}_{jt}^{k\neq\ell}$ in the second-stage regression, instead of the presence indicators $\mathbf{d}_{jt}^{k\neq\ell}$, we can only do so if $\eta_{j,t-\tau}^{\ell}$ and ϵ_{jt}^{k} are uncorrelated across locations ($k \neq \ell$). This is a drawback of identification under Assumption 2.

Suppose we are interested in a broader definition of the extensive margin,

$$\bar{y}_{jt}^{\text{ext},\ell} \equiv \mathbb{E}\left[\epsilon_{jt}^{\ell} \,|\, d_{jt}^{\ell} = 1; \mathbf{z}_{j,t-\tau}\right],$$

which does not condition on the observed location pattern outside ℓ. This definition allows us to investigate the impact of a permanent wage differential (in $\mathbf{z}_{j,t-\tau}$) through its effect on the entire grid of an MNE's potential locations. Formally, we can now evaluate $\nabla_{\mathbf{p}_{jt}} m^{\ell}(\mathbf{p}_{jt}) \cdot \nabla_{\mathbf{z}_{j,t-\tau}} \mathbf{p}_{jt}$ (a matrix), where $\mathbf{p}_{jt}$ is the vector of propensity scores. Under the restriction that $\eta_{j,t-\tau}^{\ell}$ and ϵ_{jt}^{k} are not correlated across locations ($k \neq \ell$), d_{jt}^{k} is not correlated with ϵ_{jt}^{k} because ϵ_{jt}^{ℓ} must be uncorrelated with all functions of $\mathbf{z}_{j,t-\tau}$. Then we can relax item (i) in Assumption 2 to $\mathbb{E}[\epsilon_{jt}^{\ell} \,|\, d_{jt}^{\ell} = 1, \mathbf{z}_{j,t-\tau}] = m^{\ell}(\mathbf{p}_{jt})$.

Assumption 3

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

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

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

Proposition 3 follows as a corollary to Proposition 2 (replace the scalar derivative $\partial\Delta m^{\ell}(p_{jt}^{\ell}, \mathbf{d}_{jt}^{k\neq\ell})/\partial p_{jt}^{\ell}$ with the vector $\nabla_{\mathbf{p}_{jt}} \Delta m^{\ell}(\mathbf{p}_{jt})$, and $\nabla_{\mathbf{z}_{j,t-\tau}} p_{jt}^{\ell}$ with $\nabla_{\mathbf{z}_{j,t-\tau}} \mathbf{p}_{jt}$).

Proposition 3 *If Assumption 3 holds and if* $m^{\ell}(\mathbf{p}_{jt})$ *and* $p_{jt}^{\ell}(\mathbf{z}_{j,t-\tau})$ *are continuously differentiable and have continuous distribution functions almost everywhere, then* $\mathbf{x}_{jt}^{\ell}\beta^{\ell}$ *and* $m^{\ell}(\mathbf{p}_{jt})$ *are identified up to additive constants.*

Das, Newey, and Vella (2003) establish convergence rates and asymptotic normality of similar estimators on the basis of smoothness properties of $p_{jt}^{\ell}(\mathbf{z}_{j,t-\tau})$ and $m^{\ell}(\mathbf{p}_{jt})$ (and a generalization of $\mathbf{x}_{jt}^{\ell}\beta^{\ell}$ to a function of $\mathbf{x}_{jt}^{\ell}$) for splines and power series. We use power series to approximate $p_{jt}^{\ell}(\mathbf{z}_{j,t-\tau})$ and $m^{\ell}(\mathbf{p}_{jt})$. Power series are root-$n$ asymptotic normal and can estimate smooth functionals of unknown parameters (Newey 1997). Most important for our application, the first derivative of the power series estimator is a smooth functional and hence also root-n asymptotic normal.

Estimation. We first estimate equations (5) with individual linear regressions by location. We use a third-order polynomial in wages and two additional predictors, alongside otherwise linear predictors (to break the curse of dimensionality). Second, we include the predicted propensity scores $\hat{p}^{\ell}_{jt}$ from the first stage on the second stage (6). Under Assumption 2 we approximate $m^{\ell}(p^{\ell}_{jt}, \mathbf{d}^{k \neq \ell}_{jt})$ with a third-order polynomial in p^{ℓ}_{jt}, interacted with $\mathbf{d}^{k \neq \ell}_{jt}$ (we continue to include absence indicators $(\mathbf{1}{-}\mathbf{d}_{jt})$ without interactions to both approximate $m^{\ell}(\cdot)$ and remove potential stacking bias). Under Assumption 3 we approximate $m^{\ell}(\mathbf{p}^{\ell}_{jt})$ with a third-order polynomial in $\mathbf{p}_{jt}$ (and include absence indicators $(\mathbf{1}{-}\mathbf{d}_{jt})$ among the regressors to remove potential stacking bias). We implement the second-stage estimation of (6) for $L{-}1$ locations (excluding home) by iterating SUR over the estimated disturbance covariance matrix until the estimates converge. Through constraints, we impose linear homogeneity in factor prices and symmetry of wage coefficients (see appendix B).

3.3 Wage Elasticities of Labor Demand

We use elasticities of substitution to quantify the responses of multinational labor demand y^{ℓ}_{jt} to permanent wage changes. The (constant-output) *cross-price elasticity of substitution* between factors ℓ and k is defined as $\varepsilon_{\ell k} \equiv \partial \ln y^{\ell}_{jt} / \partial \ln w^k$ and becomes

$$\varepsilon^T_{\ell k} = \frac{\psi_{\ell k} + s^{\ell} s^k}{s^{\ell}} \quad (k \neq \ell) \qquad \text{and} \qquad \varepsilon^T_{\ell \ell} = \frac{\psi_{\ell \ell} + s^{\ell}(s^{\ell} - 1)}{s^{\ell}} \tag{10}$$

for a short-run translog cost function function, where $s^{\ell} = w^{\ell} y^{\ell} / C$ is the wage bill share of the workforce at ℓ (the wage bill at location ℓ in the MNE's total wage bill) and $\psi_{\ell k} \equiv \partial s^{\ell}_{jt} / \partial \ln w^k$ is the marginal change of the wage bill share at ℓ in response to a log wage change at k. These elasticities can be calculated both for each individual MNE-j observation and in the aggregate using sample means. We will report elasticity estimates from cost function coefficients and observed mean wage bill shares.

A *permanent* change of the wage level w^k in location k is reflected in both vectors of regressors $\mathbf{x}^{\ell}_{jt}$ (with w^k_t) and $\mathbf{z}_{j,t-\tau}$ (with $w^k_{t-\tau}$). So, the response of the wage bill share s^{ℓ}_{jt} to a permanent change in $\ln w^k_t$ is

$$\psi_{\ell k} = \delta_{\ell k} + \partial \mathbb{E}\left[\epsilon^{\ell}_{jt} \,|\, \cdot, w^k_{t-\tau}\right] / \partial w^k_{t-\tau} \equiv \psi^{\text{int}}_{\ell k} + \psi^{\text{ext}}_{\ell k}. \tag{11}$$

The first term in (11) captures the labor demand response at the intensive margin $\psi^{\text{int}}_{\ell k} \equiv \partial s^{\ell}_{jt} / \partial w^k_t$. The second term in (11) is a measure of the labor demand response to a permanent change in w^k at the extensive margin $\psi^{\text{ext}}_{\ell k} \equiv \partial s^{\ell}_{jt} / \partial w^k_{t-\tau}$.

By (6), the labor demand response at the intensive margin is $\psi^{\text{int}}_{\ell k} = \delta_{\ell k}$ under any of the Assumptions 1 through 3. The labor demand response at the extensive margin, however, depends on the identifying assumption:

$$\psi^{\text{ext}}_{\ell k} = \begin{cases} \gamma^{\ell}_{w^k}\beta^{\ell}_{\Lambda}\,\Delta^{\ell}_{jt}\cdot w^{\ell}_t w^k_t/C_{jt} & \text{Assumption 1,} \\ (\partial m^{\ell}(p^{\ell}_{jt},\mathbf{d}^{k\neq\ell}_{jt})/\partial p^{\ell}_{jt})\cdot(\partial p^{\ell}_{jt}/\partial w^k_{t-\tau})\cdot w^{\ell}_t w^k_t/C_{jt} & \text{Assumption 2,} \\ \nabla_{\mathbf{p}_{jt}} m^{\ell}(\mathbf{p}_{jt})\cdot\nabla_{w^k_{t-\tau}}\mathbf{p}_{jt}\cdot w^{\ell}_t w^k_t/C_{jt} & \text{Assumption 3.} \end{cases} \tag{12}$$

We multiply by present wages w^k_t because estimation on the first stage uses w^k_t as regressors, not their logs. We divide by C_{jt}/w^{ℓ}_t to convert estimates from labor demand equation (6) back into their wage bill share equivalents because we also use $\psi^{\text{int}}_{\ell k} = \delta_{\ell k}$ at the intensive margin. Under Heckman (1979) correction (Assumption 1), $\gamma^{\ell}_{w^k}$ is the wage coefficient in the selection equation, $\beta^{\ell}_{\Lambda} \equiv \rho^{\ell\ell}_{\epsilon\eta}\sigma^{\ell}_{\epsilon}$ is the coefficient on the selectivity hazard in the outcome equation, and Δ^{ℓ}_{jt} is the first derivative of the selectivity hazard Λ^{ℓ}_{jt} (the inverse of the Mills ratio) with respect to its scalar argument, $\Delta^{\ell}_{j}(\mathbf{z}_{j,t-\tau}\gamma^{\ell}) \equiv \Lambda^{\ell}_{j}(\mathbf{z}_{j,t-\tau}\gamma^{\ell})[\Lambda^{\ell}_{j}(\mathbf{z}_{j,t-\tau}\gamma^{\ell}) - \mathbf{z}_{j,t-\tau}\gamma^{\ell}]$. Because $\Delta^{\ell}_{j}(\cdot) \in (0,1)$, the sign of the log wage effect on the wage bill at the extensive margin is the sign of the product $\gamma^{\ell}_{w^k}\beta^{\ell}_{\Lambda}$ (the coefficients on the two stages of estimation). Under polynomial series estimation, the derivatives of $m^{\ell}(\cdot)$ and p^{ℓ}_{jt} are the marginal effects on the third-order polynomials, evaluated at the sample mean.[10]

We run 200 bootstraps on the two-stage procedure to find standard errors for our elasticity estimates. Bootstrapping is advantageous because it does not require treatment of insignificant wage coefficients from the first-sage regressions in our quantification of the extensive margin. Moreover, Eakin, McMillen, and Buono (1990) show in simulations that analytic confidence intervals for elasticity estimates under normality assumptions can widely differ from bootstrapped confidence interval estimates.

4 Data and Descriptive Statistics

Our main data source is a confidential three-dimensional panel (parent-affiliate-year observations) of German MNEs at Deutsche Bundesbank (BuBa). We retain manufacturing parents and majority-owned manufacturing affiliates only. We transform the data to parent-location-year observations and combine the data with complementary information on wages and host-country characteristics from various sources.

Firm-level data. Information on foreign affiliates' turnover, employment and fixed assets stems from BuBa's MIDI database (MIcro database Direct Investment, formerly DIREK). MIDI contains outward FDI information from a legally mandated

[10] If w^{ℓ}_t is a strictly location-specific variable, equation (12) does not apply to $k = \ell$ since w^{ℓ}_t drops from a binary probit likelihood function. By our variable construction, w^{ℓ}_t is MNE j's competitors' mean factor price exposure. It is thus also MNE-specific.

Table 1: Employment at German MNEs in 2000

	HOM (1)	CEE (2)	DEV (3)	OIN (4)	WEU (5)
Employment	1,423,086[a]	245,721	332,622	319,221	394,579
Estimation sample 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

Sources: MIDI and USTAN 1996 to 2001, manufacturing MNEs and their majority-owned foreign manufacturing affiliates. 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.

annual survey, which covers the universe of German firms and households with foreign corporate holdings above minimum ownership shares and capital stock thresholds (Lipponer 2003). Individually identified outward FDI data are available for the years 1996-2001 and provide two-digit NACE 1.1 sector classifications for the parent and affiliates. We restrict our sample to majority-owned foreign affiliates because estimation of a multilocation cost function suggests the use of observations of parent firms with full managerial control and because majority ownership is insensitive to a change in the notification threshold in MIDI 1999. Assets and capital structure of every majority-owned foreign firm are reported in MIDI, including in years with zero turnover. Turnover does not distinguish within-MNE shipments from final sales but is nevertheless a proxy to affiliate production for cost function estimation.

Balance sheet and income statement information for German parent firms comes from BuBa's USTAN database, which records this information for German firms that draw a bill of exchange (for a documentation in English see Deutsche Bundesbank (1998)). The bill of exchange is a common form of payment among firms of all sizes throughout the sample period 1996-2001 (though losing some popularity thereafter), and USTAN is considered the most comprehensive source of balance sheet data for companies of all sizes outside the financial sector in Germany. The MIDI and USTAN data were linked by parent name and address in previous work (Becker, Ekholm, Jäckle, and Muendler 2005), resulting in the loss of some observations from the universe.[11]

To obtain interpretable results, we lump host countries into four *aggregate locations*: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), and WEU (Western Europe); see table 15 in the Appendix for definitions. As Table 1 shows, the four aggregate foreign locations

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

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 population 1996 and 2000 (not matched to USTAN), manufacturing MNEs and their majority-owned foreign manufacturing affiliates. Locations: Home (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe); see table 15 for definitions.

host similarly large manufacturing workforces for German manufacturing MNEs: between 250,000 and 400,000 employees. Aggregation into four foreign locations beyond home reduces the estimated cross-wage labor demand elasticity matrix to five columns and rows (with 25 elasticity estimates). Except for possibly DEV, which spans Latin America and the Asia-Pacific region (except Japan, Australia and New Zealand), aggregate locations are fairly homogeneous. Among the low-wage locations we focus on CEE, where most expansions happen. Among the 2,247 MIDI MNEs with foreign presence either in 1996 or 2000, CEE was the region where MNEs opened most new affiliates, 18.2 percent more in 2000 than in 1996, followed by DEV with 12.6 percent, OIN with 3.2 percent and WEU with 2.0 percent.

MIDI and USTAN matches are incomplete so that we do not observe parent employment for every German MNE. For comparisons, we predict total parent employment for the full sample of German manufacturing MNEs from a linear regression of parent employment on foreign employments and estimate that German manufacturing MNEs with majority-owned foreign manufacturing affiliates employ about 1.4 million German workers. Conditional on MNE presence, the largest employment per sample MNE occurs in OIN and the smallest employment in WEU.

Table 2 shows changes to the presence patterns of German MNEs between 1996 and 2000. Adjustments are infrequent. Among firms who remain MNEs in both

Table 3: MNE COUNTS OF CHANGING AFFILIATE NUMBERS

$N_{2000} - N_{1996}$	CEE (1)	DEV (2)	OIN (3)	WEU (4)	*MNE Total* (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 population 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. Locations: Home (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.

years, more than four in five with a presence in only one location abroad in 1996 keep exactly one foreign location (large numbers in row 2; large numbers sum to 100 percent for location counts 2 through 5). More than half of all MNEs who are present in only one foreign location in 1996 have a presence in only one foreign location in 2000 (small numbers in row 2; small numbers sum to 100 percent for location counts 1 through 5). In general, 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 not gradual: MNEs who exit most frequently abandon all foreign locations at once; frequencies in the first column dominate frequencies below the diagonal in every row (small numbers in column 1). There is a large number of complete withdrawals between 1996 and 2000 (477 out of 2,247 MNEs). 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 Deutsche Bundesbank's commitment to follow up on missing questionnaires.

German MNEs typically pursue a single-affiliate strategy of foreign expansions: the median number of affiliates of a German MNE per 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 observations of

MNEs in given locations exhibit no change to the number of affiliates between 1996 and 2000; 247 out of 1,259 observations of MNEs in their locations increase or decrease the number of affiliates by one. A small remainder of 153 manufacturing 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.) Together, the infrequent changes to foreign presence in Tables 2 and 3 suggest that MNEs face potentially large sunk costs of foreign presence.

Changes to the number of host countries within locations are even more infrequent 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 in given locations exhibit no change in the number of selected host countries within the location. Infrequent net changes to the number of affiliates and countries could, in principle, conceal gross changes such as changes to the country composition within a location or exit and reentry with a different affiliate. Yet only small fractions of MNEs who maintain a constant number of affiliates within a location change countries in the location. In both CEE and WEU 4.2 percent of MNEs with constant affiliate numbers between 1996 and 2000 change 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 changes beyond net changes are mostly country changes and not reentries with different affiliates. Motivated by these findings, we define the extensive margin (selection into a location) as the presence of an MNE in an aggregate location with at least one affiliate. We do not distinguish the few country changes within aggregate locations for selection estimation, but our labor demand (outcome) estimation accounts for varying country-level exposures.

We deflate parent variables with the German CPI and deflate affiliate variables with country-level CPIs (from the IMF's International Financial Statistics). CPI deflation factors are re-based to unity at year end 1998. We 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. See Appendix C for details on currency conversion.

Complementary data. Wage information is not reported in MIDI. We obtain 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). To mitigate possible workforce composition effects in our labor demand regression on wages, we use medians over sectors by foreign country. Though German wages are available from USTAN, we also take the German wages from UNIDO for comparability; we use sector wages for location selection estimation (where workforce composition behind labor cost measures is not an econometric concern) and Germany-wide sector medians for translog estimation. We conduct ro-

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 Fixed assets (t)	17.264	14.886	15.108	15.804	15.282
ln Turnover (t)	18.450	15.931	16.505	17.277	17.073
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. 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).

bustness checks using OWW wage data by occupation (Occupational Wages around the World, Freeman and Oostendorp 2001) between 1983 and 1999 and using UBS wage data for 1994, 1997, 2000 and 2003. We also obtain sector-specific German wages from the original data that underly the OWW information for Germany. We deflate and currency-convert the wages in accordance with all other variables, and transform them into annual wages. Appendix D provides further details on wage variable construction.

National accounts information for host-country regressors comes from the World Bank's World Development Indicators and the IMF's International Financial Statistics. We use CEPII bilateral trade and geographic data (*www.cepii.fr*) to compute market access to a host country as in Redding and Venables (2004), see Appendix E. To condition selection estimation on skill endowments beyond labor costs, we include the host country's percentage of high-school or higher 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%).[12]

Table 4 shows means of variables by location in the censored panel (of MNEs with presence in at least one foreign location for labor demand estimation). In our main specifications, we consider multinational labor demand during the years 1998-2001 (called t) for a sample of 1,640 MNEs and infer their location selection two years prior to production ($t - \tau$) from an uncensored sample of 3,392 MNEs. For robustness checks, we also use a single cross-section of 322 MNEs in 2000 and their location selection in 1996. The frequency of MNE presence abroad increased by two to four percentage points between 1996-99 and 1998-2001 in all locations but WEU (Western European countries) where it slightly fell in the censored panel. German MNEs spend the bulk of their wage bill (79 percent) at home. 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 smaller conditional differential could reflect MNE selection into relative high-wage countries within the low-wage region CEE.

Choice-specific variables (host country attributes) are not identified in binomial choice models such as probit for parametric selection correction. We estimate our model also in an MNE cross-section where we have no time-varying host country attributes. We therefore transform host country attributes to competitor-averages by MNE, and use competitor-average transformations in all procedures for comparability. We group MNEs into eight manufacturing sectors[13] and calculate mean host-country attributes over all competitor observations by location and sector. We take the total of competitors' foreign employments as host-country weights within the location. The wage at $t - \tau$ in CEE, for example, is the average wage paid at competitor's affiliates in CEE. In Table 4, we only take means over MNEs with presence in a given location so that the table reports CEE wages of the competitors

[12] For estimation of location selection, we also experimented with German import and export data from 2000 as controls for trade in the MNE's home sector. The import and export data were at the two-digit product level (matching NACE 1.1 two-digit sector codes) and by country of destination or origin (Fachserie 7, Reihe 7 from *destatis.de/genesis*) but did not prove to be significant predictors of location selection.

[13] The 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; manufactures not elsewhere classified.

of a German MNE with FDI in CEE.[14] German MNEs in CEE, compared to any other location, face competitors in host countries that offer the least market access, that have the smallest skill endowments, that are geographically the closest and that exhibit the smallest per-capita consumption. The CEE wages paid by competitors of MNEs in CEE are below those paid by competitors in DEV. MNEs in OIN face 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 West 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.

5 Estimation

A permanent wage differential between an MNE's home and a foreign location directly affects employment at the intensive margin through labor reallocation across existing affiliates. A permanent wage differential indirectly affects labor demand at the extensive margin by altering the likelihood of presence, which in turn changes conditional expectations of labor demand. We estimate both margins.

The effect of home wages on employment is identifiable at both margins from sector variation in a cross-section of German MNEs because individual wage-taking firms face bargained earnings schedules from sectoral agreements between unions and employers' associations (with one-year to two-year terms).[15] Time variation of home wages provides additional identification. Similarly, both time variation and variation across locations identify employment effects of foreign wages at the intensive margin. Identification of foreign wages at the extensive margin is more limited, however. Because binomial choice models (of presence or absence) cannot identify coefficients of choice-specific variables (host country attributes), foreign wage changes at the extensive margin are mainly identified over time. We obtain additional variation by considering competitor-average foreign wages which vary by MNE. To clear wage variables of workforce composition effects, we use country-wide sector medians for foreign wages. For German wages, we use sector medians in outcome (translog) estimation but sector wages in location selection estimation (where composition

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

[15]The use of sector home wages and location selectivity controls removes potential firm-level bargaining effects behind labor demand coefficients on home wages. Foreign affiliates of German MNEs are few and small, with arguably no impact on foreign wage levels.

Table 5: Sunk-cost Coefficients in Short Probit Regression

	CEE	DEV	OIN	WEU
	(1)	(2)	(3)	(4)
FDI in CEE $(t-\tau)$	2.112	-.181	-.131	-.290
	(.060)***	(.067)***	(.071)*	(.058)***
FDI in DEV $(t-\tau)$	-.169	2.200	.124	-.156
	(.069)**	(.063)***	(.070)*	(.061)**
FDI in OIN $(t-\tau)$	-.149	.146	2.274	-.140
	(.071)**	(.069)**	(.066)***	(.063)**
FDI in WEU $(t-\tau)$	-.461	-.220	-.310	1.760
	(.056)***	(.059)***	(.062)***	(.051)***
Const.	-.872	-1.241	-1.319	-.707
	(.044)***	(.049)***	(.050)***	(.042)***
Obs.	3,392	3,392	3,392	3,392

Sources: MIDI 1996 to 2001, pooled sample of manufacturing MNEs and their majority-owned foreign manufacturing affiliates with two-year selection lags ($\tau = 2$). Standard errors in parentheses: * significance at ten, ** five, *** one percent. Locations: Home (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

effects in wages are not a concern). Estimation at the intensive margin conditions on a firm's MNE status.

5.1 Location choice

We estimate binomial choices of presence in up to four foreign locations—CEE, DEV, OIN and WEU—with probit regressions for parametric selectivity correction (Assumption 1) and with series estimators of selection propensities for nonparametric correction (Assumptions 2 or 3).

Probit estimation. To have a first idea of sunk costs in location choice, Table 5 shows probit probability estimates from a short regression of MNE presence on past presence indicators across locations and a constant. Past presence between 1996 and 1999 at a given location is a highly significant predictor of MNE presence two years later in that location (and continues to be highly significant in a long regression). MNE presence indicators elsewhere serve as rudimentary controls. We consider this regression a reduced-form version of the empirical presence rule (5); long regressions that underpin location selection with additional economic regressions will corroborate the sunk cost implication that past presence predicts about 70 percent of the propensity of future presence.

Table 6: Sunk Entry and Exit Costs in Probability Terms

	CEE	DEV	OIN	WEU
	(1)	(2)	(3)	(4)
Sunk entry cost: γ_N	.872***	1.241***	1.319***	.707***
	(.044)	(.049)	(.050)	(.042)
Sunk exit cost: γ_X	1.240***	.959***	.954***	1.053***
	(.291)	(.225)	(.224)	(.247)
Hysteresis band: $(\gamma_N + \gamma_X)$	2.112***	2.200***	2.274***	1.760***
	(.060)	(.063)	(.066)	(.051)
Marginal effect of hysteresis band	.704***	.710***	.714***	.621***
	(.015)	(.016)	(.017)	(.014)

Sources: MIDI 1996 to 2001, 3,392 pooled observations of manufacturing MNEs and their majority-owned foreign manufacturing affiliates with two-year selection lags. Estimates are probit coefficients from Table 5. Significance levels from χ^2 tests. Standard errors in parentheses: * significance at ten, ** five, *** one percent. Locations: Home (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

The reduced-form estimates provide a summary view of sunk costs in probability terms. Recall that the sunk cost part of location selection (5) can be represented with

$$F^{\ell}_{j,t-\tau} = \gamma^{\ell}_N - (\gamma^{\ell}_X + \gamma^{\ell}_N)\, d^{\ell}_{j,t-\tau},$$

where γ_N are sunk entry costs, γ^{ℓ}_X sunk exit costs, and $(\gamma^{\ell}_X + \gamma^{\ell}_N)$ is also called the hysteresis band. Table 6 shows the decomposition result, based on estimates of coefficients along the diagonal and the constant in Table 5. For the entry and exit cost decomposition involves the estimate of the constant, 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.

Past presence increases the likelihood of future presence in a given location by more than seventy percent in all but WEU, where the marginal effect predicts a more than sixty percent increase. Long probit regressions confirm these magnitudes. The total, however, hides the differential impact of entry and exit costs. Entry cost are the largest in the distant low-income and high-income locations DEV and OIN, and dominate exit costs there. Conversely, entry costs are the 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.

Indicators for past FDI presence may not exclusively capture sunk costs but also firm heterogeneity. In long regressions, we look into the black box behind rule (5) and

Table 7: Marginal Effects in Long Probit Regressions

Predictors $(t-\tau)$	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. × Home sector wage	-.0007	-.005	-.015	-.020
	(.005)	(.004)	(.004)***	(.008)***
FDI in CEE × Comp.s' wages CEE	.054	-.060	-.093	.090
	(.066)	(.057)	(.050)*	(.083)
FDI in OIN × 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 ($\tau = 2$). Standard errors in parentheses: * significance at ten, ** five, *** one percent. Further regressors (not significantly different from zero at five percent level in any location): Competitors' wages DEV and WEU and their interactions with FDI presence in DEV and WEU, Competitors' hosts ln Market access, Indic. of Headquarters West Germany, ln Equity, Parent profits/equity, Competitors' hosts skill shares, Competitors' hosts distance, Competitors' hosts ln Consumption per capita. Without wage-presence interactions, past presence has a marginal effect of .779 (standard error .022) in CEE, .671 (.027) in DEV, .713 (.026) in OIN, and .747 (.020) in WEU. Locations: CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

include firm-level predictors as well as competitor-average host country attributes. Table 7 presents the marginal effects for the full list of covariates.[16] Among the firm-level predictors, we include interactions between past presence indicators and wages to capture the co-determining effect of wage differentials and an MNE's past presence at a location.

In the long regressions of Table 7, past presence elsewhere (off the diagonal) loses predictive power. But past presence for the region itself continues to be a statistically significant and salient predictor of presence (except for CEE because of the wage interaction). 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—similar to the marginal effects in the short regression (last row in Table 6). These probability effects of past presence confirm the importance of the hysteresis band. The MNE's number of host countries in the past also significantly raises the likelihood of presence. German MNEs with large home employment, low parent debt, and a high capital-labor ratio at the parent firm are significantly more likely to be present at most or all foreign locations within two years.

Time and, by construction, competitor variation identifies wage effects. The home wage has the expected positive sign in all regressions and is a significant predictor for presence in OIN and WEU, both by itself and in its interaction with past presence. The negative coefficients on the home wage interaction with past presence suggest that wage differentials matter less for the location decision of MNEs that already own an affiliate in the region. With German wages partly controlling for the wage differential between the foreign location and the home sector, several foreign wages are statistically insignificant predictors of location choice. Insignificant coefficients of host wages are common in the literature on location choice (e.g. Devereux and Griffith (1998) for U.S., Head and Mayer (2004) for Japanese, and Buch, Kleinert, Lipponer, and Toubal (2005) for German MNEs). For estimation of the cross-elasticity of labor demand at the extensive margin, however, only the coefficient on the German wage matters (because the extensive margin is only defined for foreign labor demand). Bootstraps over both estimation stages will show even for the statistically weak wage prediction of location selection into CEE that, weighted with the strong labor demand effects of CEE selection, home wage levels significantly affect the elasticities of labor substitution at the extensive margin.

Further covariates (not reported) include competitors' wages in OIN and WEU and their interactions with past presence in DEEV and WEU, competitors' hosts'

[16] For continuous variables, marginal effects are $\gamma^{\ell *} = \partial\Phi(\cdot)/\partial z_{j,t-\tau} = \phi(\cdot)\gamma^{\ell}$; for indicator variables, marginal effects are the differences in $\Phi(\cdot)$ between setting the indicator to 1 or 0 (evaluated at the sample mean $\bar{\mathbf{z}}_{j,t-\tau}$, and the variance-covariance matrix estimator being transformed with the delta method). Sample size drops from 3,392 to 2,414 mainly because of missing information from parent balance sheets.

market access, competitors' hosts' skill shares, competitors' hosts' distance, competitors' hosts' per-capita consumption, an indicator of parents' headquarters in West Germany, equity, and parent profits/equity. None of those covariates is significant at the five-percent level in any location. 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 significant predictors of location choice; here we leave them out.

Nonparametric propensity score approximations. We estimate the propensity score of location choice with a third-order polynomial in wages, market access, and the count of an MNE's past host countries, alongside the same linear predictors as for probit estimation. The predicted propensity scores 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.

Table 8 reports coefficient estimates by location. Marginal effects are close to those in the probit regressions. Estimates of the hysteresis band along the diagonal of past presence indicators 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—again close to the marginal effects in the short regression (last row in Table 6). 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 wage-presence interaction countervails the high marginal effects of past presence.

We present 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 important predictors for cross-elasticities of labor substitution at the extensive margin. Using UNIDO wages, 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 remain significant predictors under nonparametric estimation, excepting the host country count variable. Similarly, insignificant parent-level covariates remain insignificant.

5.2 Translog estimation with selectivity correction

We proceed to estimation of the short-run translog cost function and include predicted selection hazards from probit estimation as regressors in the equation system (parametric selectivity correction, Assumption 1). Alternatively, we include pre-

Table 8: MARGINAL EFFECTS IN NONPARAMETRIC PROBABILITY MODEL

Predictors $(t-\tau)$	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. × Home sector wage	.001	.006	-.010	-.004
	(.003)	(.004)*	(.003)***	(.004)
FDI in OIN × 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 ($\tau = 2$). 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' hosts ln Market access, ln Count of host countries, Indic. of Headquarters West Germany, ln Equity, Parent profits/equity, Competitors' hosts skill shares, Competitors' hosts distance, Competitors' hosts 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.

dicted propensity scores from nonparametric selection estimation (Assumptions 2 or 3).

Translog cost function estimation. Table 9 presents estimates of translog cost function equations for 1,640 stacked MNE observations between 1998 and 2001. (We lose observations mainly because of missing wage information at affiliate locations.) Coefficient estimates are from iterated seemingly unrelated regressions of transformed wage bill shares on their translog predictors for four out of five locations, excluding home. For the regression, wage bill shares and covariates are scaled by observation-specific cost-wage ratios to remove upper truncation. 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 9 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.

Equation estimates in the upper panel of Table 9 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 translog cost function regressors predict the bulk of labor demand variation across locations, with R^2 regression fits ranging between .92 and .97 for all equations. Regression fits are similar under parametric and nonparametric selectivity correction. Overall, we consider the significance of selectivity correction terms strong evidence for the importance of the extensive margin.

Tests for parametric selectivity correction. We test whether Assumption 1 for parametric selection correction is satisfied in our context. There are two cases: (a) the part of the selection shock that correlates with labor demand shocks is an MNE-specific disturbance and does not vary by location, and (b) the labor-demand related part of the selection shock varies by location but is independent of labor demand shocks in other locations. We test the two cases in turn. Tests fail to reject case (a), but they do reject case (b). We consider the assumptions of case (a) both economically plausible and statistically acceptable.

Consider (a) MNE-specific selection shocks whose labor demand related part does not vary by location. This case implies that the covariance between selection disturbances is the same for any pair of locations $k \neq \ell$. We obtain estimates of those covariances from multivariate probit estimation of simultaneous selection into the four foreign locations. In the cross section of MNEs in 2000 with multivariate

Table 9: TRANSLOG COST PARAMETER ESTIMATES

Labor cost shares in: (transformed)	CEE (1)	DEV (2)	OIN (3)	WEU (4)
Parametric Selectivity Correction (Assumption 1)				
ln *Wages*				
HOM	.020	-.002	.078	.183
	(.001)***	(.0008)**	(.004)***	(.005)***
CEE	-.008	-.001	-.003	-.008
	(.0008)***	(.0002)***	(.0004)***	(.0005)***
DEV	-.001	.001	-.002	.004
	(.0003)***	(.0008)	(.0004)***	(.0006)***
OIN	-.003	-.002	-.112	.039
	(.00007)***	(.00007)***	(.003)***	(.001)***
WEU	-.008	.004	.039	-.219
	(.0001)***	(.0001)***	(.001)***	(.004)***
Selectivity hazard	81.487	32.872	33.468	92.618
	(15.830)***	(17.751)*	(12.462)***	(16.618)***
R^2	.945	.950	.966	.932
Nonparametric Selectivity Correction (Assumption 2)				
ln *Wages*				
HOM	.023	.0003	.075	.149
	(.001)***	(.001)	(.005)***	(.006)***
CEE	-.008	-.003	-.003	-.009
	(.0008)***	(.0004)***	(.0005)***	(.0006)***
DEV	-.003	.002	-.002	.003
	(.0004)***	(.0009)**	(.0005)***	(.0007)***
OIN	-.003	-.002	-.109	.040
	(.0005)***	(.0005)***	(.005)***	(.002)***
WEU	-.009	.003	.040	-.183
	(.0006)***	(.0007)***	(.002)***	(.006)***
Series terms				
χ^2 tests (p-value)	517.4 (.000)	376.0 (.000)	117.8 (.000)	198.9 (.000)
R^2	.954	.955	.965	.926

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages). 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 (omitted), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

probit regressors from 1996, we fail to reject the joint equality of six correlation coefficients between the four equations with a χ^2 test statistic of 4.63 (p value .592). Under a restriction on the selection disturbance, another implication of case (a) is that, if an MNE is neither present in all locations nor absent from all locations, the choices of presence and absence must be consistent with a location-independent MNE-specific selection shock for all locations. We calculate the regression prediction for all MNEs that are not omnipresent, pick the maximum probit prediction among all locations of absence and the minimum probit prediction among all locations of presence thus stacking the cards of the test against case (a). Although 191 out of 1,941 observations show the wrong pattern, a one-sided t-test of the null hypothesis that absence and presence predictions are reversed is rejected overwhelmingly with a t statistic of 77.4 (zero p value).

Second, consider (b) location-varying selection shocks that are independent of labor demand shocks in other locations. This assumption implies that labor demand residuals from one location must have no correlation with labor demand residuals from any other location. We reject this hypothesis for three out of six pairs of the four location residuals with p values below .01, for two pairs with p values below .1, but fail to reject zero correlation in one remaining case (CEE-OIN).

While case (b) is rejected, there is no evidence against case (a) where selection disturbances correlate with labor demand shocks only through an MNE-specific shock but not through location-specific shocks. Note that cross-location correlations of labor demand errors are not evidence against case (a) because MNE-specific selection shocks themselves induce a correlation between the labor demand disturbances across locations. As discussed before, case (a) is plausible in economic terms. Suppose selection disturbances include both host country-specific parts such as, for example, surprising changes to profit repatriation policies and include MNE-specific parts. 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-relevant host-country shocks so that cost function disturbances are unrelated to the entry-relevant policy shocks. Given supportive test results and the economic plausibility of case (a), we regard estimation under parametric selectivity correction (Assumption 1) a relevant benchmark.

Elasticities of multinational labor substitution. Table 10 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 (Germany) in a sample of MNEs, which are observable to the statistician only when active in the home location. The standard errors are from 200 bootstraps over the two estimation steps of parametric selectivity corrected translog estimates (Assumption 1). One margin at a time is set to zero to isolate the effect at the other margin. Cross-price elasticities are affine transformations of translog coefficients (equation (10)). While

Table 10: 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). 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, Table 9). 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).

log wage effects on wage bill shares are additive in translog estimation at the intensive and the extensive margin (equation (6)), cross-wage substitution elasticities are not.

Own-wage elasticities along the diagonal—for both the intensive and the extensive margins—are uniformly negative, 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. Note that 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 because those constraints are not well defined at the observation level—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 is consistent with two main 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 in absolute terms. Second, CEE workers tend to be less productive than German workers, which is reflected in the translog cost function coefficients.

The extensive margin is a noticeable component of adjustment, beyond its crucial role in correcting cost function estimates for location selectivity bias. 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 7), the strong significance of the selection effect on labor demand on the second stage in CEE (selectivity hazard coefficient in Table 9) turns home wage effects into significant predictors of employment substitution at the extensive margin. Beyond the marginal wage coefficients from two-step estimation, observed wage bill shares provide information for elasticity estimation and thus contribute to the significance of elasticity estimates.

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 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. In CEE, a one-percent increase in the German home wage is also associated with a .8 percent increase in MNE employment at the extensive margin, if no adjustment occurs at the intensive margin.

We also add the intensive and extensive wage effects on wage bills and compute the total home wage elasticities of foreign labor demand. We find highly significant estimates for the total elasticities at three locations: .61 in CEE, 2.51 in OIN and 2.45 in WEU (significantly different from zero at the one-percent level). Our 200 bootstraps allow us to test whether the elasticities at the intensive margin are significantly different from the total elasticities. We reject their 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

Table 11: Foreign-Wage Elasticities of Home Employment

Home employment change (%)	Wage change (1%) in HOM (1)	CEE (2)	DEV (3)	OIN (4)	WEU (5)	Obs. (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, OWW 98-01	-.477	.051	-.002	.209	.219	1,458
	(.053)***	(.010)***	(.005)	(.030)***	(.037)***	
Ass. 1, UBS 98-01	-.434	.013	.008	.078	.336	1,614
	(.056)***	(.006)**	(.011)	(.031)**	(.038)***	
Ass. 2, UNIDO 98-01	-.533	.055	.014	.146	.319	1,640
	(.048)***	(.006)***	(.006)**	(.026)***	(.032)***	
Ass. 3, 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 wages). 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).

by a positive cross-wage elasticity between OIN and WEU (1.14) at the extensive margin.

5.3 Specification comparisons

To assess the robustness of our estimates, we compare several further specifications and report the first rows of the cross-wage elasticity matrices (foreign wage effects on home employment) in Table 11, and the first columns separately by intensive and extensive margin in Tables 12 and 13 (home wage effects on foreign employment).

Foreign-wage elasticities of home employment are robust across specifications (Table 11). 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

Table 12: Home-Wage Elasticities at the Intensive Margin

	Home wage change (1%), by regression specification						
	Stacking						Omnipr.
	UNIDO	UNIDO	UBS	OWW	UNIDO	UNIDO	UNIDO
	98-01	00	98-01	98-01	98-01	98-01	98-01
Emplmt.	Ass. 1	Ass. 1	Ass. 1	Ass. 1	Ass. 2	Ass. 3	Ass. 1
chg. (%)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CEE	1.596	1.810	1.366	.603	1.707	1.648	3.535
	(.218)***	(.748)**	(.247)***	(.272)**	(.215)***	(.226)***	(4.062)
DEV	.651	1.534	-.147	.322	.807	.880	-.444
	(.466)	(1.004)	(.480)	(.430)	(.323)**	(.397)**	(1.072)
OIN	2.328	2.573	3.540	.979	2.255	2.235	1.938
	(.432)***	(.888)***	(.516)***	(.399)**	(.376)***	(.363)***	(.482)***
WEU	2.214	1.860	2.087	1.826	1.951	1.915	2.851
	(.224)***	(.407)***	(.353)***	(.197)***	(.191)***	(.205)***	(.494)***
Obs.	1,640	322	1,458	1,614	1,640	1,640	93

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages). 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: HOM (Germany), CEE (Central and Eastern Europe), DEV (Developing countries), OIN (Overseas Industrialized countries), WEU (Western Europe).

main source of identification at the intensive margin. OWW and UBS wage data in the third and fourth row result in smaller estimation samples and perhaps introduce attenuation bias for some coefficients (the UBS wage data are particularly sketchy for CEE). Coefficient estimates are nevertheless similar across wage data. Nonparametric estimation does not yield statistically different estimates, neither under Assumption 2 nor 3, excepting DEV. The sample of 93 omnipresent MNEs between 1996 and 2001 is small but results in significant outcome estimates on the second stage (we predict selectivity hazards from first-stage regressions on the full sample); the magnitude of coefficient estimates, when significant, is considerably larger than for the stacked samples, suggesting that foreign employment at omnipresent MNEs responds more elastically to home wages. 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 (Table 12) are robust too. Estimates on our benchmark sample (now in the first column) conform closely to several other specifications. In fact, the comments on the rows of Table 11 above apply also to the columns of Table 12, except only that the coefficient estimates for the sample of omnipresent MNEs now closely resemble those from other

Table 13: Home-Wage Elasticities at the Extensive Margin

	Home wage change (1%), by regression specification						
	Stacking						Omnipr.
	UNIDO	UNIDO	UBS	OWW	UNIDO	UNIDO	UNIDO
	98-01	00	98-01	98-01	98-01	98-01	98-01
Emplmt.	Ass. 1	Ass. 1	Ass. 1	Ass. 1	Ass. 2	Ass. 3	Ass. 1
chg. (%)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CEE	.795	.838	.395	.524	.869	-.040	.643
	(.201)***	(.232)***	(.380)	(.197)***	(3.282)	(9.586)	(.300)**
DEV	.772	.572	.975	.626	-9.719	3.941	.592
	(.162)***	(.252)**	(.298)***	(.892)	(8.133)	(17.680)	(.503)
OIN	.960	1.116	1.431	1.160	.833	-4.249	.345
	(.340)***	(.392)***	(.845)*	(.625)*	(3.669)	(7.373)	(.331)
WEU	1.016	1.183	1.561	1.808	1.527	-2.457	.719
	(.171)***	(.301)***	(.372)***	(.504)***	(1.999)	(3.141)	(.096)***
Obs.	1,640	322	1,458	1,614	1,640	1,640	93

Sources: MIDI and USTAN 1996 to 2001 (UNIDO wages). 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).

specifications.

At the extensive margin, home-wage elasticities of foreign employment (Table 13) are (highly) significant in the parametric specifications (columns 1 through 4), for all wage data and in the year 2000 cross section (with UNIDO wages). Coefficient magnitudes vary slightly more across specifications than they do at the intensive margin. 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 outlier predictions, for which the first-stage probability model results in propensity scores outside the zero-one range. Nonparametric estimates for the extensive margin (columns 5 and 6 of Table 13) are not statistically different from zero but similar in magnitude when plausible (column 5, excepting DEV). Although the inclusion of nonparametric series terms in translog estimation yields more precise estimates of intensive margin coefficients (Tables 11 and 12) by approximating disturbance components beyond the parametric selectivity hazard, the series terms do not seem to provide a precise estimate of the extensive margin itself.We nevertheless view the similarity between parametric and plausible nonparametric estimates as 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

Table 14: Counterfactual Employment Effects of a One-percent Reduction in the Home-Foreign Wage Gap

	Permanent wage gap reduction by one percent between Home and			
Employment effect	CEE	DEV	OIN	WEU
at the intensive margin on	(1)	(2)	(3)	(4)
Home[a]	728	161	2,141	5,143
	(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). Point estimates from parametric selectivity correction (Assumption 1, Table 10) 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).

this selected sample expands to foreign locations more frequently.

Taken together, our results confirm the statistical plausibility of the benchmark estimates from parametric selectivity correction (Assumption 1). Several tests for the validity of Assumption 1 fail to reject the identifying hypothesis that selection shocks correlate with labor demand shocks only through an MNE-specific error but not through location-specific errors. Nonparametric estimation yields very similar and highly significant elasticity estimates at the intensive margin. At the extensive margin, the benchmark estimates from parametric selectivity correction are highly significant but nonparametric estimates fail to attain significance. In short, the benchmark estimates from parametric selectivity correction are statistically robust. We now turn to the economic implications of our estimates for multinational labor substitution.

5.4 Counterfactual Evaluation

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 14 give answers to these questions.

We use the home-wage elasticities of foreign labor demand and the foreign-wage elasticities of home labor demand from our selectivity corrected translog benchmark estimates for the 1998-2001 MNE sample (Table 10). These estimates reflect the employment responses at the mean MNE (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 700 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 wag 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 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 (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 (96.1) percent of that in Germany so that an increase in WEU wages by 14% (2%) to cut the gap in about half would attract 70,000 (4,000) 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 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.

6 Conclusion

While the public discourse over outsourcing seems to have settled on the idea that multinational enterprises (MNEs) substitute jobs at home for foreign employment, 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 cost function 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 foreign affiliates. We derive conditions for common Heckman (1979) selectivity corrections, location by location, and for nonparametric identification. Cost function estimation, however, conditions on MNE output. The empirical exercise thus leaves aside the counterfactual question how the market share and size of an MNE would differ if its access to foreign locations were limited in spite of global product market competition. This matter is part of our ongoing research.

Empirical evidence on German manufacturing MNEs shows that firms change multinational presence only infrequently and hardly alter their number of affiliates within regions. These infrequent changes to multinational presence at the extensive margin give rise to rare but salient labor demand effects in response to permanent wage differentials across locations. With every percentage increase in Central and Eastern European wages, German manufacturing MNEs are found to allocate 700 MNE jobs to Germany. 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 salient impact on multinational labor substitution.

Appendix

A A Model of the MNE

We think of an MNE's choice of multinational production and sales as a two-stage decision problem. At moment $t-\tau$ (i.e. τ periods prior to production and sales), MNE j determines at which locations to produce and faces uncertainty over other MNEs' future output $\mathbf{q}_{i\neq j,t}$, input prices $\mathbf{w}_t$, and its own realized output $\mathbf{q}_{jt}$. With its location choice, the MNE also chooses its optimal capital stock vector $\mathbf{k}_{jt}$ across L locations.

On the second stage at time t, all uncertainty is removed and MNE j chooses output $\mathbf{q}_{jt}$ given its cost function (or, by duality, optimal factor employment given its production function). The optimal quantity choice $\mathbf{q}^*_{jt}$ at time t can be characterized with the first-order conditions

$$p^\ell(\mathbf{q}_{i\neq j}, \mathbf{q}^*_{jt}) \left(1 - 1/\varepsilon^{q\ell}_{jt}\right) \leq \frac{\partial C_{jt}(\mathbf{q}^*_{jt}; \mathbf{k}_{jt}, \mathbf{w}_t)}{\partial q^\ell_{jt}} \qquad (\ell = 1, \ldots, L), \tag{A1}$$

where $p^\ell(\cdot)$ is the price of a good from location ℓ as a function of competitors' and own worldwide output, and $\varepsilon^{q\ell}_{jt}$ is the elasticity of demand for q^ℓ_{jt} with respect to price p^ℓ. By the Kuhn-Tucker theorem, $\mathbf{q}^*_{jt}=0$ if inequality holds. So, even if MNE j is present at location ℓ, it may find it optimal to produce $q^\ell_{jt}=0$ once factor price and competitors' output are revealed.

On the first stage, MNE j's linear programming problem can be characterized by the rules for FDI at locations $\ell = 1, \ldots, L$

$$d^\ell_{jt} = \mathbf{1}\Big(\mathbb{E}_{j,t-\tau}\Big[\Pi_{jt}(q^{\ell,*}_{jt}) - \Pi_{jt}(q^\ell_{jt}=0)\Big|\, \mathbf{z}_{j,t-\tau}\Big] - F^\ell_{j,t-\tau} + \eta^\ell_{j,t-\tau} > 0\Big), \tag{A2}$$

where $F^\ell_{j,t-\tau}$ denotes MNE j's relevant fixed costs for presence at location ℓ and $\eta^\ell_{j,t-\tau}$ is an MNE-specific disturbance. Expectations depend on MNE j's information set $\mathbf{z}_{j,t-\tau}$. MNE j's linear programming problem on the first stage involves the simultaneous evaluation of (A2) for each location ℓ given the 2^{L-1} possible combinations of outputs at all remaining locations $L-1$.

For its location choice on the first stage, an MNE j maximizes its expected profits $\mathbb{E}_{j,t-\tau}[\Pi_{jt}]$ where expectations are conditional on the MNE's information set in period $t-\tau$. The MNE can produce the vector of outputs $\mathbf{q}_{jt} = (q^1_{jt}, \ldots, q^L_{jt})'$ at L locations ($\ell = 1, \ldots, L$). So, future expected profits are

$$\mathbb{E}_{j,t-\tau}\left[\,\mathbf{p}(\mathbf{q}_{i\neq j,t}, \mathbf{q}_{jt})' \cdot \mathbf{q}_{jt} - C_{jt}(\mathbf{q}_{jt}; \mathbf{k}_{jt}, \mathbf{w}_t)\,\right]. \tag{A3}$$

The estimated presence rule (5) in the text follows using expected profits (A3) in criterion (A2).

B Multiproduct translog cost function

Consider the short-run multiproduct translog function with quasi-fixed capital:[17]

$$
\begin{aligned}
\ln C_{jt} \;=\;& \varphi + \sum_{m=1}^{L} \varphi_m^0 \ln q_{jt}^m + \sum_{\ell=1}^{L} \alpha_\ell \ln w_t^\ell + \sum_{m=1}^{L}\sum_{\ell=1}^{L} \mu_{\ell m} \ln q_{jt}^m \ln w_t^\ell \\
&+ \frac{1}{2}\sum_{m=1}^{L}\sum_{\ell=1}^{L} \varphi_{\ell m}^1 \ln q_{jt}^m \ln q_{jt}^\ell + \frac{1}{2}\sum_{m=1}^{L}\sum_{\ell=1}^{L} \delta_{\ell m} \ln w_t^m \ln w_t^\ell \\
&+ \sum_{m=1}^{L} \zeta_m^0 \ln k_{jt}^m + \sum_{m=1}^{L}\sum_{\ell=1}^{L} \zeta_{\ell m}^{11} \ln k_{jt}^m \ln q_{jt}^\ell \qquad \text{(B1)} \\
&+ \sum_{m=1}^{L}\sum_{\ell=1}^{L} \kappa_{\ell m} \ln k_{jt}^m \ln w_t^\ell + \frac{1}{2}\sum_{m=1}^{L}\sum_{\ell=1}^{L} \zeta_{\ell m}^1 \ln k_{jt}^m \ln k_{jt}^\ell .
\end{aligned}
$$

By Shepard's (1953) 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

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

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 equation (6).

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 $\sum_{\ell=1}^{L} \mu_{\ell m} = \sum_{\ell=1}^{L} \kappa_{\ell m} = \sum_{\ell=1}^{L} \delta_{\ell m} = \sum_{\ell=1}^{L} \delta_{m\ell} = 0$ for all m. We impose those restrictions on estimation but do not constrain estimates of factor price coefficients at the extensive margin. We do not impose any returns-to-scale restrictions.

C Currency conversion and deflation

We convert all economic data of foreign affiliates into euro (EUR) and deflate them. 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. We apply the following deflation and currency conversion method to all financial variables. Deutschmark (DEM) figures are converted into euro figures at

[17] Slaughter (2000) adds $\ln(k/q)$ terms to a version of (B1). 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 $\ln(k/q) = \ln k - \ln q$.

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

We use the foreign countries' CPIs (Consumer Price Indices from the IMF's International Financial Statistics) to deflate the figures. Whenever a country's CPI is not available from IFS but the main currency used in that country is issued in some other country, we use the CPI of the currency-issuing country. The CPI deflation factors for all countries are rebased to unity at year-end 1998. For the UBS wage data, we first translate U.S. dollars into Euros and then proceeded as detailed above. Parent-level and sector-level domestic variables are transformed into December 1998 Euros using the German CPI.

D Wages

We base our estimation on 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. We deflate the wages with the country-level CPI (standardized to unity in December 1998) and convert the foreign currency to EUR at the December 1998 exchange rate. To mitigate possible workforce composition effects in our labor demand regression on wages, we use the sector median wage by country (and lose sectoral wage variation also for Germany) in the outcome estimation. We use sectoral UNIDO wages for Germany in selection estimation because workforce composition behind labor cost measures is not an econometric concern for location choice. The UNIDO data cover 109 countries and result in the largest overlap with MIDI observations on German MNEs for estimation.

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 on German MNEs 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 deflate the wages with the country-level CPI (standardized to unity in December 1998) and convert the foreign currency to EUR at the December 1998 exchange rate. 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, deflate them with the German CPI (standardized to unity in December 1998), and multiply them by twelve to arrive at annual earnings for cost function estimation. 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.

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

Whenever foreign price deflators are missing or period-average exchange rate information is incomplete for purchasing-power parity oriented wage conversion, we use current exchange rates and the German price deflator.

E Market access

We construct market access measures following Redding and Venables (2004). We obtain bilateral trade data for 1996 through 2001 and geographic information on

Table 15: LOCATION DEFINITIONS

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
DEV	Developing countries including Turkey, Russia and Central Asian economies as well as dominions of Western European countries and the USA

country pairs from CEPII (*www.cepii.fr*). After filling in missing imports to B from A with exports information from country A to B, we drop all exports information and set exports from A to B equal to B's imports from A. We adopt this procedure because we consider imports, whenever available, more reliably measured than exports.

Our regression specification for an unbalanced panel of country pairs by year is

$$\ln X_{ij} = \alpha_i x_i + \beta_j m_j + \delta \ln d_{ij} + \mu\, b_{ij} + \epsilon_{ij},$$

where X_{ij} denotes country i's aggregate exports in USD (+1) to country j, x_i an exporter country dummy, m_j an importer country dummy, d_{ij} the geographical distance between country i and j, and b_{ij} a dummy variable indicating a common border. We compute market access A_i to country i as

$$A_i = \exp\{\beta_i\, m_i\} \left(.67\sqrt{area_i/\pi}\right)^{\delta/2} + \textstyle\sum_{j \neq i} \exp\{\beta_j\, m_j\}\, (d_{ij})^{\delta} \exp\{\mu\, b_{ij}\}.$$

This is measure MA(3) in Redding and Venables (2004).

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