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Foreign Direct Investment and Wages: Does the Level of Ownership Matter?

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Foreign Direct Investment and Wages: Does the Level of Ownership Matter?*

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Abstract

This paper examines the relationship between foreign equity participation and average wages at the plant level. I show that using a binary measure for foreign ownership, as is the traditional practice in the literature, leads to biased estimates of the foreign ownership wage premium, compared to the use of a continuous measure if the true relationship is linear. Using nonparametric and semi-parametric techniques I find this is the case: the relationship between the level of foreign ownership and average wages is better approximated as linear rather than binary. I find that a ten percentage point increase in foreign equity participation is associated with an approximately 4% increase in the average wage of non-production workers. These results are the first to show that the wage premium due to foreign ownership varies with the level of foreign ownership in a continuous manner.

JEL Classification: C33; F23; J31
Keywords: Foreign Direct Investment; Wages; Censoring; Dynamic Panel Data

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1 Introduction

A large body of trade research is devoted to understanding the host-country effects of foreign direct investment (FDI), which has increased dramatically in recent decades. One of the key effects studied extensively in this regard is the impact that multinational activity has on average wages at plants subject to foreign acquisition. It is now well understood that affiliates of multinational companies pay higher wages compared to their domestic counterparts even after controlling for sectoral, regional, and plant-level characteristics.[1] However, existing studies provide a range of estimates for the average wage effect of multinational status from 1 percent to 70 percent. Despite the similarity in the methodology and data sets employed in these studies, it remains to be understood why we observe such a large range of estimates at seemingly similar multinational plants and what the precise wage effects of multinational activity are. As Girma and Gorg (2007) note, even when controlling for observable and time invariant unobservable characteristics, there remains a fundamental problem in identifying the performance differences that is attributable to multinationality *per se*. The current study identifies the causes of such divergent estimates and documents the causal impact of foreign ownership on wages using methodology that sidesteps earlier limitations.

Existing studies often estimate a firm-level wage model with a binary variable that indicates multinational status. Using binary variables in the estimation embodies the assumption that all multinational affiliates are identical with respect to the wages they pay. Since an indicator variable for ownership status censors the information on what share of the affiliate equity is controlled by the multinational parent, it is unable to capture any variation in the wages due to different levels of foreign equity participation. If the level of control that multinationals exercise at their affiliates affects the wages they pay, then estimation with a binary variable will fail to capture the heterogeneity in the wage premium across multinational affiliates. Moreover, an econometric issue arises if the wage premium varies with foreign equity participation. Rigobon and Stoker (2009) show that the least squares estimator is prone to severe bias when there are several regressors and a binary variable is used in place of a continuous regressor. Estimates also become susceptible to the level of thresholds in defining the multinational status

[1] See, for example, Aitken et al. (1996), Feenstra and Hanson (1997), Doms and Jensen (1998), Figini and Gorg (1999), Taylor and Driffield (2005), Lipsey and Sjoholm (2006), Almeida (2007), Heyman et al. (2007), Girma and Gorg (2007), and Arnold and Javorcik (2009). See Table A1 in Almeida (2007) for a summary of the literature on the multinational wage premium using firm level data and the premium estimates.

of an affiliate. Thus, using a censored foreign ownership variable, as is the common practice in the literature, would lead to inconsistent estimates of the wage premium when wages vary with the level of foreign equity participation.

This study identifies the heterogeneity in the multinational wage premium that arises due to the level of foreign equity participation using a unique data set from Turkey. The distinguishing feature of these data is the observation of continuous levels of foreign ownership at the plant level with a considerable degree of ownership distribution across the plants. I build on the results by Rigobon and Stoker (2009) to show that using censored regressors may lead to severe bias not only in ordinary least squares estimation, but also in fixed effects estimation. More specifically, when the true relationship between foreign equity participation and average wages is linear, using a binary variable instead of a continuous regressor leads to inconsistent estimates of the wage premium even if plant level individual effects are accounted for. By artificially creating different thresholds for the foreign ownership variable, I illustrate the "variability" of wage premia across different definitions and the biases that ensue. Up to 14 percent of the wage premium attributed to a foreign owner may come from different levels of foreign equity participation even after controlling for plant level effects.

Two main results come out of the empirical analysis, which uses the census of Turkish manufacturing plants over the period 1993-2001.[2] First, using nonparametric and semiparametric regressions, I demonstrate that there is essentially a linear and increasing relationship between the level of foreign ownership and average wages. This monotonic relationship holds more strongly for non-production workers than production workers. Second, I find that a significant wage premium exists only for non-production workers when I produce estimates of the premium that control for plant level effects and the endogeneity of foreign ownership. I address the endogeneity of multinational activity by generating instruments from the panel data at hand in a generalized method of moments framework, which allows me to accommodate a large set of assumptions on the estimated wage model. My results indicate that a 10 percentage point increase in foreign equity participation is associated with a 4 percent increase in the average wage of non-production workers, and that the level of foreign ownership does not affect the wages of production workers. Therefore, there is a significant degree of heterogeneity

[2]Three earlier studies report estimates from matched employer-employee data in addition to firm-level estimates (see Heyman et al. (2007) for Sweden, and Martins (2004) and Almeida (2007) for Portugal). These estimates focus on whether foreign firms pay higher wages to identical workers. While it is desirable to have such data to control for worker heterogeneity, such employer-employee data do not exist for Turkey.

in the wage premia at multinational affiliates that comes from different levels of foreign equity participation. Moreover, the finding that there is no significant premium for production workers is novel in the literature.

Existing literature has identified how the level of foreign ownership is related to certain aspects of the firm, which may have an impact on average wages. Takii and Ramstetter (2005) find that higher foreign equity participation is associated with higher levels of productivity in Indonesian manufacturing. A similar finding is documented in the case of Venezuela by Aitken and Harrison (1999). This could arise because a majority foreign ownership share might be required for bringing in technologies from the parent firm, which in turn may lead to a high wage premium (Lipsey and Sjoholm, 2006). In a similar vein, Barbosa and Louri (2002) argue that a foreign partner will demand higher ownership in case of profitable affiliates and large intangible assets to be transferred. Indeed, Budd et al. (2005) find that the degree of multinational ownership appears to condition the degree of intrafirm profit sharing, and that affiliate wages are positively correlated with both parent and affiliate profits.

Although some existing studies consider the impact of the level of foreign ownership on the wage premium, there is no consensus in the literature on the subject. On the one hand, Martins (2004) finds no higher wage premia for firms that exhibit a stronger degree of foreign control in Portugal. On the other hand, Lipsey and Sjoholm (2006) and Aitken et al. (1996) find that majority-owned foreign plants pay higher wages for skilled workers in Indonesia and Venezuela, respectively. However, these studies do not address the endogeneity of foreign ownership explicitly. Few studies, notably Heyman et al. (2007), Girma and Gorg (2007), and Arnold and Javorcik (2009), present estimates of the wage premium that tackle the issue of endogeneity by using matching techniques. Hence, the current study is the first in the literature to identify systematic heterogeneity in the wage premium due to different levels of foreign ownership while accounting for endogeneity explicitly. I find that foreign equity participation impacts average wages at every level and my results are not driven by those multinationals achieving majority control at their affiliates.

The rest of the paper is structured as follows. The next section discusses the empirical strategy of earlier studies and builds on the results from Rigobon and Stoker (2009) to demonstrate the problems with the use of binary regressors in the panel data context. Section 3 introduces the data to be used in the analysis. Section 4 presents my empirical strategy to test the implications of Section 2 and to identify the relationship between the level of foreign ownership and average wages. Section 5 includes my empirical results and presents a set of robustness checks. Concluding remarks appear in Section 6.

2 Censoring the Level of Foreign Ownership

The equity share that a multinational controls at an affiliate is often unobserved in plant-level data. When it is observed, the common practice is to designate a certain threshold and define a plant as "foreign-owned" if the multinational's equity participation exceeds that threshold.[3] In this section, I discuss three issues. First, I describe how the common practice of using different thresholds to define foreign ownership can hide the heterogeneity in the wage premium. Second, building on Rigobon and Stoker (2009), I derive the bias in the fixed effects estimate of the wage premium that arises from censoring a continuous regressor in a single variable regression. Lastly, I extend this result to the multivariate case and discuss how censoring the level of foreign ownership disaffects the estimation of the wage premium.

Assume that the true empirical model that links wages to foreign ownership at the plant level is given by:

$$w_{i,t} = \alpha_i + m(x_{i,t}) + \gamma' \mathbf{y}_{i,t} + \varepsilon_{i,t}, \quad i = 1, ..., N; \, t = 1, ..., T \tag{1}$$

where $w_{i,t}$ represents the potential wage, $x_{i,t} \in [0, 100]$ denotes foreign equity participation in percentages at plant i at time t, $m(.)$ is a function that relates $x_{i,t}$ to wages, α_i is a time-invariant plant effect, $\mathbf{y}_{i,t}$ is a vector of plant-level controls, and $\varepsilon_{i,t}$ is White noise. If the level of foreign ownership affects wages linearly, then the true model becomes:

$$w_{i,t} = \alpha_i + \beta x_{i,t} + \gamma' \mathbf{y}_{i,t} + \varepsilon_{i,t}, \quad i = 1, ..., N; \, t = 1, ..., T \tag{2}$$

I confirm in the later sections that the estimated relationship between foreign equity participation and wages is indeed linear for the present study. I am interested in the wage premium due to the level of multinational activity, which is captured by β in (2). The inclusion of α_i in (2) enables the identification of β from within-plant variation in

[3] In national and international accounting standards, FDI is typically defined as involving an equity stake of 10 percent or more at the plant level (Razin and Sadka, 2007), although different countries follow different recording practices. For instance, Sweden uses the 50% cut-off in defining foreign ownership (Heyman et al., 2007). While researchers typically use this cut-off to define majority control, it has been noted by the finance literature that shareholders can achieve effective control in many cases by holding a block that is much smaller than 50% of the firm (Razin and Sadka, 2007).

foreign control, thus sidestepping problems that might arise from selection of high-wage plants by multinationals.

Earlier studies estimate a wage premium by using a censored version of the foreign ownership variable mostly because their data prevented them from observing $x_{i,t}$ in its continuous nature. Specifically, they estimate:

$$w_{i,t} = a_i + bF_{i,t} + c'\mathbf{y}_{i,t} + \epsilon_{i,t}, \;\; i = 1, ..., N; \; t = 1, ..., T \tag{3}$$

where $F_{i,t}$ is a binary variable indicating foreign ownership, defined on a threshold, ϕ:

$$F_{i,t} = 1[x_{i,t} > \phi] \tag{4}$$

Implicit in this practice is the assumption that all foreign plants are identical. When this is not the case, Figure 1 depicts how censoring the level of foreign ownership hides the heterogeneity in the wage premium and leads to different estimates depending on the threshold.[4] In the figure, all domestic firms are assumed to pay the same wage, w_1, while wages are increasing in the level of foreign ownership for multinationals, as depicted by the function $m(x)$. Assume that we estimate this relationship with an equation such as (3), and we set $\phi = 0\%$. The variable of interest, $\hat{b}$, will capture an effect illustrated by l_1 in the figure, with every multinational predicted to pay w_2. As l_1 simply captures an average effect, it overstates the wage premium for multinationals with less than 50 percent ownership and understates it for those above this level. If we instead set $\phi = 50\%$, then $\hat{b}$ captures an effect illustrated by l_2, at which all multinationals are predicted to pay w_3. In this case, l_2 overestimates the wage premium for most multinationals and provides a higher estimate than l_1. Hence, censoring not only hides the heterogeneity in the wage premium due to the level of foreign ownership, but it also results in confounded estimates due to lack of knowledge on $m(x)$.[5]

Rigobon and Stoker (2009) derive the bias from using censored regressors for the OLS (ordinary least squares) estimator, and I build on their results for the case of 0-1 censoring as in (4). I show here that their results can be readily extended to the FE (fixed effects, or within-group) estimator. In order to motivate the result, I start the analysis with a single regressor. Let the true model be given by (2), excluding the vector of controls $\mathbf{y}_{i,t}$. The fixed effects transformation eliminates α_i from (2) and yields a single variable model in deviations from individual means:

[4]Figure 1 is hypothetical and intended for demonstrative purposes only.

[5]Note that if there was no heterogeneity in the wage premium, then $\hat{b}$ would return the same estimate independent of the value of ϕ and accurately capture the return to being a multinational.

$$w_{i,t} - \bar{w}_i = (x_{i,t} - \bar{x}_i)\beta + (\varepsilon_{i,t} - \bar{\varepsilon}_i) \tag{5}$$

where $\bar{w}_i = T^{-1}\sum_{t=1}^{T} w_{i,t}$, and $\bar{x}_i$ and $\bar{\varepsilon}_i$ are defined similarly. The FE estimator, which is unbiased in finite samples, is given by:

$$\hat{\beta}_{FE} = \frac{\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{i,t} - \bar{x}_i)(w_{i,t} - \bar{w}_i)}{\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{i,t} - \bar{x}_i)^2}$$

I am interested in the asymptotic bias that arises when one estimates the following model instead:

$$w_{i,t} = a_i + bF_{i,t} + \epsilon_{i,t}, \;\; i = 1, ..., N; \, t = 1, ..., T \tag{6}$$

where the $F_{i,t}$ is defined as above. The coefficient of interest is estimated by:

$$\hat{b}_{FE} = \frac{\sum_{i=1}^{N}\sum_{t=1}^{T}(F_{i,t} - \bar{F}_i)(w_{i,t} - \bar{w}_i)}{\sum_{i=1}^{N}\sum_{t=1}^{T}(F_{i,t} - \bar{F}_i)^2}$$

The bias that I am going to characterize is given by $plim\, \hat{b}_{FE} - \beta$, which will clearly be affected by the threshold ϕ. To see this formally, recall that $\hat{b}_{FE}$ is identical to the estimator obtained by an OLS estimation of the dummy variable model:

$$w_{i,t} = \sum_{j=1}^{N} a_j d_{i,j} + bF_{i,t} + \epsilon_{i,t} \tag{7}$$

where $d_{i,j} = 1$ if $i = j$ and 0 elsewhere. Following Rigobon and Stoker (2009), the probability limits of the OLS estimators of (7) are given by:[6]

$$\begin{aligned}
plim\, \hat{a}_{i,FE} &= E[w_{i,t}|F_{i,t} = 0, \alpha_i] = \alpha_i + \beta E[x_{i,t}|F_{i,t} = 0, \alpha_i] \\
plim\, \hat{b}_{FE} &= E[w_{i,t}|F_{i,t} = 1, \alpha_i] - E[a_i|F_{i,t} = 1, \alpha_i] \\
&= E[w_{i,t}|F_{i,t} = 1, \alpha_i] - E[w_i|F_{i,t} = 0, \alpha_i] \\
&= \alpha_i + \beta E[x_{i,t}|F_{i,t} = 1, \alpha_i] - \alpha_i - \beta E[x_{i,t}|F_{i,t} = 0, \alpha_i] \\
&= \beta \{E[x_{i,t}|F_{i,t} = 1, \alpha_i] - E[x_{i,t}|F_{i,t} = 0, \alpha_i]\}
\end{aligned}$$

The FE estimator $\hat{b}_{FE}$ measures β up to a positive scalar as in the OLS case, but differently, this scalar is now determined by the expectations conditional on α_i. The

[6] The difference here from Rigobon and Stoker (2009) is the conditional expectations, since the true data generating process (DGP) is now given by the single variable version of (2) with time-invariant individual effects instead of a cross-sectional DGP. Remember that the interpretation of β comes from the conditional expectation on the structural equation (2) even though one uses the censored version of (5) or (7) in practice to estimate the parameters of the model.

bias is:

$$plim\,\hat{b}_{FE} - \beta = \beta\,\{E[x_{i,t}|F_{i,t} = 1,\, \alpha_i] - E[x_{i,t}|F_{i,t} = 0,\, \alpha_i] - 1\}$$

What does this result tell us? If one is merely interested in whether foreign ownership causes a positive or negative wage premium, then using a censored regressor will provide a consistent answer as to the direction of this association. However, if the interest is in the size of the premium, then $\hat{b}_{FE}$ provides an estimate that is confounded by the difference $E[x_{i,t}|F_{i,t} = 1,\, \alpha_i] - E[x_{i,t}|F_{i,t} = 0,\, \alpha_i]$. This *within* difference depends not only on ϕ, but also on the conditional distribution of the uncensored variable $x_{i,\,t}$. For instance, if foreign owners acquire higher equity stakes at plants that are larger in size or that operate in certain industries, then we would expect the within difference to be larger in such plants and industries. Thus, the extent of the heterogeneity in foreign ownership directly impacts the wage premium estimate and 0-1 censoring might lead to misestimates by hiding this information.

In practice, one is typically interested in the parameters of a multivariate model, which calls into question the transmission of bias among the regressors. Assume that the true model is given by (2) in which the vector $\mathbf{y}_{i,\,t}$ consists of a single control $y_{i,\,t}$. The censored model is:

$$w_{i,\,t} = a_i + bF_{i,\,t} + cy_{i,\,t} + \epsilon_{i,\,t}, \;\; i = 1,\,...,\,N;\; t = 1,\,...,\,T \tag{8}$$

The FE estimator of b is again identical to the estimator obtained by OLS estimation of the dummy variable model:

$$w_{i,\,t} = \sum_{j=1}^{N} a_j d_{i,\,j} + bF_{i,\,t} + cy_{i,\,t} + \epsilon_{i,\,t} \tag{9}$$

Following Rigobon and Stoker (2009), denote the residual of $w_{i,\,t}$ regressed on $F_{i,\,t}$ as: $\Delta w_{i,\,t} = w_{i,\,t} - (1 - F_{i,\,t})\bar{w}_{0,\,t} - F_{i,\,t}\bar{w}_{1,\,t}$; where $\bar{w}_{1,\,t} = \sum_{i=1}^{N}\sum_{t=1}^{T} F_{i,\,t}w_{i,\,t}/\sum_{i=1}^{N}\sum_{t=1}^{T} F_{i,\,t}$ is the average of $w_{i,\,t}$ for $F_{i,\,t} = 1$, and $\bar{w}_{0,\,t} = \sum_{i=1}^{N}\sum_{t=1}^{T}(1 - F_{i,\,t})w_{i,\,t}/\sum_{i=1}^{N}\sum_{t=1}^{T}(1 - F_{i,\,t})$. Applying the same transformation to both sides of (2), one gets:

$$\Delta w_{i,\,t} = \beta\Delta x_{i,\,t} + \gamma\Delta y_{i,\,t} + \Delta\varepsilon_{i,\,t} \tag{10}$$

If one applies this transformation to the model in (9), both the censored variable $F_{i,\,t}$ and the individual dummies $d_{i,\,j}$ are removed, which yields the estimation equation:

$$\Delta w_{i,\,t} = c\Delta y_{i,\,t} + v_{i,\,t} \tag{11}$$

Rigobon and Stoker (2009) note that the bias in $\hat{c}$ of (8) is the same as that of (11), which arises due to the omission of $\Delta x_{i,t}$ from (10). The standard omitted variable bias formula then yields $plim\, \hat{c}_{FE} = \gamma + \beta\eta \equiv c$, where η is defined by:

$$\eta = \frac{Cov(\Delta y_{i,t},\, \Delta x_{i,t})}{Var(\Delta y_{i,t})} = \frac{(1-p)Cov(y_{i,t},\, x_{i,t}|F_{i,t}=1,\, \alpha_i^*) + pCov(y_{i,t},\, x_{i,t}|F_{i,t}=0,\, \alpha_i^*)}{(1-p)Var(y_{i,t}|F_{i,t}=1,\, \alpha_i^*) + pVar(y_{i,t}|F_{i,t}=0,\, \alpha_i^*)}$$

and p is the probability that $F_{i,t} = 1$. Again, the difference in the current result from that of Rigobon and Stoker (2009) for the OLS case is that the covariances and variances are now conditioned on individual effects, α_i^*, where the linear projection of $x_{i,t}$ on the additional regressor is expressed as: $x_{i,t} = \alpha_i^* + \eta y_{i,t} + r_{i,t}$.

Hence, the parameter η, which measures how within-deviations of foreign equity participation are proxied by the within-deviations of the additional regressor, determines the size of the bias in $\hat{c}$. As Rigobon and Stoker (2009) note, it is impossible to assess the bias in terms of size and direction if one has no information regarding the within-variation of $x_{i,t}$. The probability limits for the other coefficients in (8) are given by:

$$\begin{aligned}
plim\, \hat{a}_{i,FE} &= E[w_{i,t}|F_{i,t}=0,\, \alpha_i] - cE[y_{i,t}|F_{i,t}=0,\, \alpha_i] \\
&= \alpha_i + \beta E[x_{i,t}|F_{i,t}=0,\, \alpha_i] + (\gamma - c)E[y_{i,t}|F_{i,t}=0,\, \alpha_i] \\
&= \alpha_i + \beta\left[E[x_{i,t}|F_{i,t}=0,\, \alpha_i] - \eta E[y_{i,t}|F_{i,t}=0,\, \alpha_i]\right]
\end{aligned}$$

$$\begin{aligned}
plim\, \hat{b}_{FE} &= E[w_{i,t}|F_{i,t}=1,\, \alpha_i] - E[w_{i,t}|F_{i,t}=0,\, \alpha_i] + cE[y_{i,t}|F_{i,t}=0,\, \alpha_i] - cE[y_{i,t}|F_{i,t}=1,\, \alpha_i] \\
&= \alpha_i + \beta E[x_{i,t}|F_{i,t}=1,\, \alpha_i] + \gamma E[y_{i,t}|F_{i,t}=1,\, \alpha_i] \\
&\quad -\alpha_i - \beta E[x_{i,t}|F_{i,t}=0,\, \alpha_i] - \gamma E[y_{i,t}|F_{i,t}=0,\, \alpha_i] \\
&\quad -c\left\{E[y_{i,t}|F_{i,t}=1,\, \alpha_i] - E[y_{i,t}|F_{i,t}=0,\, \alpha_i]\right\} \\
&= \beta\left[E[x_{i,t}|F_{i,t}=1,\, \alpha_i] - E[x_{i,t}|F_{i,t}=0,\, \alpha_i]\right. \\
&\quad \left.-\eta\left\{E[y_{i,t}|F_{i,t}=1,\, \alpha_i] - E[y_{i,t}|F_{i,t}=0,\, \alpha_i]\right\}\right]
\end{aligned}$$

The bias in $\hat{b}_{FE}$ thus depends on two extra terms compared to the single regressor case: how the additional regressor covaries with x, and the distribution of the additional regressor conditional on censoring and α_i. With additional regressors in the picture, it is possible to have a case where $\hat{b}_{FE}$ may actually have the *wrong* sign.[7] Hence, with 0-1 censoring, it is possible to end up not only with a biased estimate of the wage premium, but also with the wrong sign on it.

[7] This will be the case whenever we have: $\frac{E[x_{i,t}|F_{i,t}=1,\alpha_i]-E[x_{i,t}|F_{i,t}=0,\alpha_i]}{E[y_{i,t}|F_{i,t}=1,\alpha_i]-E[y_{i,t}|F_{i,t}=0,\alpha_i]} < \eta$.

3 Panel Data on Turkish Manufacturing

Data on the Turkish manufacturing industry come from the Industrial Analysis Database by the Turkish Statistical Office (TurkStat), which covers all manufacturing plants in Turkey with more than ten employees, including plants controlled by foreign investors. For this study, I focus on the period 1993-2001. The inclusion of plant identification codes enables me to construct a panel and follow the plants over time. The total number of manufacturing plants varied between 10,567 in 1993 and 11,311 in 2001 (see Table A2). The percentage of foreign plants in the sample, defined as plants that have at least some level of foreign ownership, increased from 2.85 percent to 3.88 percent over the same period. The measure of foreign ownership in this study is the percentage of subscribed equity owned by the foreign investor, which varies between 0 and 100 percent. The average foreign equity participation at plants owned partially or fully by foreigners increased from 58.78 percent in 1993 to 64.33 percent in 2001.

Figure 2 depicts the distribution of foreign ownership shares for all plant-year observations for the subset of foreign plants in the sample. There is a substantial degree of heterogeneity in how much control multinational firms exercise. While most foreign plants seem to be majority owned, there is a significant number of plant-year observations with multinationals owning less than 50 percent of the plant's equity. Moreover, one sees the full range of ownership shares with sizable densities in each bin of the distribution. Similar patterns can be seen when I reproduce Figure 2 for different industries or plant sizes (results not reported here). Informed by the analytical results in the previous section, I expect this pattern in the level of foreign ownership to bias estimates of the wage premium in a censored regression.

In addition to foreign ownership, the database contains yearly information on employment, inputs, output, value added, wages and compensation, sales, inventories, additions to fixed assets, energy use, sector, and location. Plant size is measured as the total number of paid workers at a plant in any given year. I observe the number of production and non-production workers and total payments to each group in the database. In all of the analyses, total yearly wages as reported by the plants are used in the calculation of the average plant wage and the average wage for production and non-production workers, excluding any additional benefits and compensation.[8]

[8]Number of paid workers are reported for production and non-production workers four times during a given year (in February, May, August, and November) and the average of these four observations constitutes the average number of workers at the plant in a given year (i.e. the plant size).

A frequently mentioned source of possible selection bias is acquisitions of high-wage domestic plants by multinational firms, also known as cherry picking (see Lipsey and Sjoholm (2006) and Almeida (2007)). It could be the case that foreign plants acquire domestic establishments that are already highly productive and large in size and that therefore pay higher wages in general. Such selection bias would distort the results of the empirical investigation if plant effects are not controlled for. Figure 3 provides the average yearly wage for plants which experienced a takeover in the sample period by type of ownership and compares these values to the average wage in the overall sample.

Figure 3 reveals that plants which experienced a takeover during the period 1993-2001 were paying much higher wages to their workers compared to the plants in the overall sample. This holds for such plants regardless of whether they were under foreign or domestic ownership, which provides evidence to the oft-mentioned selection bias of high-wage plants by foreigners. In this case, least squares estimates will tend to capture the difference in levels between the traditionally high wage firms, which are most likely to be acquired, and the traditionally low wage firms that will almost always stay under domestic control. However, one can also see from Figure 3 that wages were higher at plants that experienced a takeover when they were under foreign ownership. This suggests that foreign ownership per se might have an impact on the average wage, even though the estimated premium after controlling for the individual firm effect is likely to be much smaller than least squares estimates.

4 Empirical Methodology

Two empirical findings characterize the activity of multinationals in Turkey with respect to the level of control they exercise and the plants they acquire. First, foreign investors choose to own any percentage of subscribed capital (equity) when they engage in FDI, allowing them to exercise various degrees of control at the acquired plant. Second, regardless of the equity share they eventually own, they target domestic plants which already pay wages that are much higher than the average. In this section, I outline a three-step empirical strategy to analyze the link between foreign ownership and wages in light of these two regularities. I first describe how the predictions of Section 2 on censoring are tested and then turn to provide estimates of the foreign ownership premium that control for plant-level effects and endogeneity.

4.1 Defining Different Thresholds

Observing foreign equity participation at the plant level allows me to define multinational status using different thresholds. In order to analyze how these different thresholds affect the wage premium, I estimate the following censored equation:

$$ln\, w_{ijt} = \beta_0 + \beta_1 FDIPlant_{ijt} + \alpha' \mathbf{X}_{ijt} + Sector + Region + Time + \varepsilon_{ijt} \qquad (12)$$

where $FDIPlant_{ijt} = 1[x_{ijt} > \phi]$ indicates multinational status, x_{ijt} is foreign equity participation and varies between 0 and 100 percent, ϕ is the threshold level, and i, j, and t index plant, sector, and year, respectively. In equation (12), w_{ijt} is the average yearly plant wage and $\mathbf{X}_{ijt}$ is a vector of plant-specific characteristics such as size and skill intensity. Sector dummy variables at the two digit level of the ISIC Rev. 2, regional dummy variables, which classify each plant belonging to one of the seven geographical regions in Turkey, and time dummy variables control for sector, region and year specific wage effects, and ε_{ijt} is a random plant-specific error component. In all my specifications, I estimate the equation of interest for three dependent variables: the average plant wage, the average wage for production workers, and the average wage for non-production workers.

I estimate equation (12) by OLS and FE using four possible values of ϕ that are arbitrarily chosen: 0%, 15%, 30%, and 50%. The goal of this exercise is to demonstrate the bias in OLS and FE estimations that arises from using different thresholds in the definition of a multinational plant. Varying estimates of β_1 due to the threshold level ϕ would indicate that the multinational wage premium depends on this arbitrary definition of multinational status. In light of the analytical results in section 2, this would suggest that the level of foreign equity participation is innately tied to average wages. If this were not so, i.e. the level of foreign equity participation does not affect average wages, then we would see identical estimates and statistical (in)significance of β_1 regardless of the threshold level. This counterfactual case would correspond to the absence of heterogeneity in the foreign ownership wage premium.

4.2 Nonparametric and Semiparametric Analysis

In my second round of estimations, I examine whether the true relationship between foreign equity participation and wages is linear. I first estimate this relationship nonparametrically using the locally weighted scatterplot smoothing (Lowess) estimator of

Cleveland (1979). Consider a regression of wages on foreign equity share, given by the model:

$$w_i = m(x_i) + \varepsilon_i, \quad i = 1, ..., N \tag{13}$$

where the error term ε_i is i.i.d. Lowess is a standard local regression estimator, whereby one lets $m(x_i)$ be linear in the neighborhood of a data point x so that $m(x_i) = m + \beta(x_i - x)$. Cleveland (1979) suggested that one minimize:

$$\sum_{i=1}^{N} \{w_i - m - \beta(x_i - x)\}^2 K\left(\frac{x_i - x}{h}\right) \tag{14}$$

with respect to m and β, where $K(.)$ is a kernel weighting function. This can be achieved by performing a weighted least squares regression of w_i against $z_i' = (1, (x_i - x))$ with weights $K_i^{1/2}$ (Pagan and Ullah, 1999). The weighted least squares regression estimates for each observation i are then used to predict the value of the dependent variable to trace out the non-parametric relationship between w and x. For implementing Lowess, I use the tricubic kernel as my weighting function, which places less weight on points near the end of the sample, and I use a bandwidth of 0.8, which uses eighty percent of the sample for each regression.[9] Despite its computational intensity, Lowess is preferable over kernel regression as it uses a variable bandwidth, robustifies against outliers, and uses a local polynomial estimator to minimize boundary problems (Cameron and Trivedi, 2005).

I implement Lowess in two different ways. The first set of Lowess regressions is run on the pooled cross-section sample of plant-year observations using average plant wage and foreign equity participation. In the second set of Lowess regressions, I include plant level fixed effects to the model in (13). Accordingly, I transform my data into within-plant deviations before estimating the non-parametric model, which allows me to control for plant-specific effects. This means that the weighted least squares estimates are identified from the within-plant variation in each local regression. Hence, I am able to identify whether changes in the level of ownership at a multinational plant over time affect the level of wages at the same plant or not.

One can question whether the identified relationship by the nonparametric analysis is driven by some omitted variables. In order to overcome this concern, I next turn to a semiparametric analysis where additional controls enter the true model parametrically

[9] I also experimented with a bandwidth of 0.5 for both my nonparametric and semiparametric estimates, which left my results unchanged.

and are additively separable from the nonparametric component. Consider the partially linear model:

$$w_i = m(x_i) + \alpha' \mathbf{X}_i + \varepsilon_i, \quad i = 1, ..., N \tag{15}$$

where $\mathbf{X}_i$ is a vector of plant characteristics. I implement the difference-based semiparametric estimator of Yatchew (1997), whereby $m(.)$ is assumed to have a bounded first derivative. Yatchew (1997) suggests ordering the data such that $x_1 < x_2 < ... < x_N$ and taking the first difference of (15). The transformed equation is then estimable by ordinary least squares. First-differencing equation (15) allows inference to be carried on α' as if there were no nonparametric component in the model. But once α' is estimated, a variety of nonparametric techniques could be applied to estimate $m(.)$ as if α' were known (Lokshin, 2006), that is, after constructing the differences $w_i - \hat{\alpha}' \mathbf{X}_i$. In my estimations, the nonlinear function $m(.)$ is estimated by the Lowess procedure outlined earlier, using a bandwidth of 0.8. Additionally, a significance test on x_i can be carried out, which tests the null hypothesis that the regression function has the known parametric form $g(x, \delta) + \alpha' \mathbf{X}_i$, where δ is an unknown parameter, against the alternative semiparametric form $m(x_i) + \alpha' \mathbf{X}_i$, where $m(.)$ is unknown. Lokshin (2006) provides details on the test.

4.3 Estimating the Foreign Equity Participation Premium

If there is evidence that censoring x_i returns biased estimates and that the true relationship is linear, then I can expect the regressions with continuous observations to provide more accurate estimates of the foreign ownership wage premium. In this subsection, I focus on quantifying the impact of foreign equity participation on average wages. For this purpose, I estimate the premium by running a set of regressions on the subset of plants that have been under multinational control at any point in the sample period. In this framework, I can test whether increases in foreign equity share translate into higher wages at the plant level.

Two considerations are in place here. First, cherry-picking of high paying domestic firms by foreign investors and the presence of unobservable firm characteristics require the inclusion of plant-level effects to the econometric specification. Second, the assumption that foreign equity participation is independent of the idiosyncratic error term can be easily violated. While it is relatively easy to handle endogeneity that arises from unobserved heterogeneity, it is much harder to handle dynamic endogeneity whereby current and past levels of wages may affect the level of foreign ownership. In addition, endogeneity bias will arise if the level of foreign ownership responds simultaneously to idiosyncratic shocks and in the case of measurement error. This naturally calls for an instrumental variable estimation; yet, it is extremely difficult (if not impossible) to come up with a valid instrument in such plant level studies.

At this point, I take advantage of the panel data at hand to use exogenous regressors in other time periods to instrument for endogenous regressors in the current time period. Consider the dynamic model:

$$ln\, w_{ijt} = \gamma ln\, w_{ij,t-1} + \beta_1 FEP_{ijt} + \alpha' \mathbf{X}_{ijt} + \delta_i + \varepsilon_{ijt}, \;\; t = 2, ..., T \tag{16}$$

where δ_i denote time independent plant-level effects and we assume foreign equity participation, FEP_{ijt}, to be endogenous. It is assumed that $|\gamma| < 1$ and ε_{ijt} are serially uncorrelated. In order to tackle the endogeneity problem, one can first-difference the model in (16) to purge δ_i, which in addition renders lagged values of $ln\, w_{ijt}$ and x_{ijt} to be valid instruments in the transformed equation. Consistent and efficient estimation can then be achieved by GMM estimators that use all available lags at each period as instruments for the equations in first differences (Arellano and Bond, 1991). Blundell and Bond (1998) extend the Arellano-Bond estimator to include more instruments that are available by assuming that first differences of instrumenting variables are uncorrelated with the fixed effects, which greatly improves efficiency and reduces the finite sample bias. However, the estimator can easily generate a large number of instruments given the availability of lags and additional moment conditions, which will lead to an overfit of the endogenous variables that tends to distort inference in finite samples.[10] In order to guard against problems due to a large number of instruments, I estimate (16) both using all available lags (and differences) as instruments and with a restricted set of instruments (to two most immediate lags).

I implement the "system GMM" estimator of Blundell and Bond (1998) in a two-step procedure and apply the finite-sample correction of Windmeijer (2005) to the standard errors. Traditionally, researchers using these GMM estimators have focused on results for the one-step estimator, partly because simulation studies suggested very modest efficiency gains from using the two-step version (Bond, 2002). The two-step estimator also tends to return standard errors that are severely downward biased when the number

[10] The problem arises because a high number of instruments means a poorly estimated optimal weighting matrix in the GMM estimator. See Roodman (2008) for a discussion of how 'instrument proliferation' can lead to serious problems when implementing these GMM estimators.

of instruments is large. However, Windmeijer (2005) finds that the two-step efficient GMM estimator with the corrected variance estimate leads to more accurate inference compared to the one-step estimator. For this reason, I report estimates of the two-step procedure with the Windmeijer correction, but I also conducted the estimation with the one-step estimator as a robustness check. The results for the one-step estimator are very similar to the results reported here and available upon request in an additional appendix.

5 Results

5.1 Estimates with Different Thresholds

In my first set of regressions, I estimate equation (12) using a binary variable that specifies whether a plant is classified as foreign (i.e. $FDIPlant$ takes on the value of unity) depending on the level of foreign equity participation.[11] Table 1 documents the differences in the estimates of the wage premium when various thresholds are used. The threshold value used to define $FDIPlant$ is given in rows (a)-(d), while columns (1)-(3) present OLS estimates and columns (4)-(6) present FE estimates for the three wage variables of interest. For example, the figure in row (b) and column (1) indicates that the OLS estimate of the average plant wage premium to multinational status is 51 percent when a plant is defined as foreign if it has at least 15 percent foreign equity participation. Controlling for plant-level effects, however, reduces the estimate of the premium to 11 percent in row (b) and column (4) when the same threshold is applied.[12] It is immediate from this discrepancy that foreign investors acquire plants that already pay high wages, justifying the motivation to focus on a wage model with plant-specific effects.

[11]In each regression, I control for a set of plant-level characteristics. These are: log plant size (as measured by the total number of employees); skill intensity (given by the ratio of skilled workforce to total plant size); ratio of production workers to total plant size; log value added per worker (data on value added provided by TurkStat); log electricity used or log inputs; sector, year, and region dummies. Sector and region dummies are replaced by plant-level effects for FE regressions. I also estimated all reported specifications controlling for log inputs instead of log electricity and my results do not change. The full set of results for the OLS and FE regressions with various thresholds, including the estimates for the controls and regression diagnostics, can be found in Tables A3 and A4 in the Appendix.

[12]If foreign investors acquire plants which already pay higher wages than the rest of the domestic plants, then we should expect to see a modest wage premium to becoming multinational. This result is consistent with the results by Lipsey and Sjoholm (2006), Almeida (2007), and Heyman et al. (2007), who find a lower premium when they control for plant level effects. Almeida (2007) shows that foreign acquisitions have small effects, typically less than 2%, on average wages at the acquired firms when "cherry-picking" is taken into account.

Estimates from Table 1 indicate that the wage premium typically increases as the threshold level ϕ gets higher. This holds true of all the OLS estimates, for which the discrepancies between the estimated premia are greater across various thresholds. The average plant wage premium is estimated to be 48 percent (row (a), column (1)) when there are no thresholds, while it is estimated to be 57 percent (row (d), column (1)) when $\phi = 50\%$. This implies that 9 percent of the average plant wage premium is purely attributable to using different thresholds. When I repeat the same exercise for production and non-production workers, I see similar discrepancies between the estimated premia. The estimated premium ranges from 17 to 22 percent for production workers (column (2)) and from 37 to 45 percent for non-production workers (column (3)), suggesting a greater degree of heterogeneity in the premium for the latter group of workers.

The problems with inference on a censored variable become more apparent in the FE estimates of Table 1, columns (4)-(6). While the discrepancies between the premium estimates for different thresholds are smaller, column (5) shows that the statistical significance of the estimate can be affected by the threshold value. In column (5), the wage premium to production workers is consistently positive, yet it is significant only when $\phi = 15\%$ (row (b)). Moreover, while the OLS estimates demonstrate a higher premium when the threshold increases, the FE estimates in columns (5) and (6) do not display such monotonicity. Similar to OLS results, however, there is greater heterogeneity in the wage premium of non-production workers even after controlling for plant level effects. Column (6) indicates that the premium estimate for this group ranges from 18 to 25 percent. Hence, Table 1 shows that using different thresholds yields inconclusive evidence on whether there really exists a wage premium at foreign plants for all groups of workers, and even if so, how large this premium is.

As a further test of how different definitions of a foreign plant affect average plant wages, I divide the sample of plants in the data into four categories depending on the percentage of equity owned by the foreign investor. I assign a value of one to a plant that has foreign equity participation from the range of intervals that I specify and run a regression where I include these intervals simultaneously as independent variables.[13] The heterogeneity in the wage premium due to the level of foreign control is more pronounced in this set of regressions, reported in Table 2. Column (1) indicates that the average wage premium at a plant with at least 50 percent foreign equity participation

[13]The intervals are 0-15%, 16-30%, 31-50%, and 51-100%.

is 58 percent, while it is 42 percent for a plant with foreign equity participation in the interval 31-50 percent, and 32 percent for a plant in the 15-30 percent interval. Controlling for plant-level effects, the estimated wage premia are 15 percent for plants with at least 50 percent foreign equity, and around 8 percent for other foreign plants (column (4)), which suggests that obtaining majority control creates an impact.

However, when I run the same regression for production and non-production workers separately, I see that the effect of equity participation can be nonmonotonic. In column (6), the estimated wage premia for non-production workers for the intervals 15-30 percent, 31-50 percent, and 51-100 percent are, 20 percent, 14 percent, and 28 percent, respectively. Hence, even after controlling for plant-level effects, up to 14 percent of the estimated wage premium can be explained by different levels of foreign equity participation. Consistent with earlier findings, column (5) shows that whether or not there is a wage premium for production workers is affected by the definition of multinational status. I find that there exists a premium (around 7 percent) for this group of workers only at plants that have at least 50 percent foreign equity participation.

These results indicate that the methodology followed in classifying a plant as foreign may significantly impact the estimated effect of foreign ownership on average wages. Censoring the foreign ownership variable in an arbitrary way hides the heterogeneity in the wage premium due to the level of foreign equity participation. Moreover, this heterogeneity may exist only for a certain group of workers, and such information will be lost when econometric analysis is carried out with binary data.

5.2 Nonparametric and Semiparametric Estimates

The results from the previous section suggest a monotonic and positive relationship between foreign equity participation and average wages, however there is also some evidence indicating nonlinearities. In order to see the true shape of the relationship, Lowess plots of equation (13) are presented in Figures 4 and 5, which use the subset of plants that have been under foreign ownership at some point in the sample period. Figure 4 depicts the relationship between foreign equity participation at the plant level versus (log) average wages in the pooled sample. Panel (a), which shows the relationship for the average plant wage, indicates an upward sloping Lowess plot line that is almost exactly the same as the linear fit. In panels (b) and (c), a similar relationship is observed for the (log) average wage of production workers and non-production workers, respectively. In all of the panels, the nonparametric fit displays an upward trend. One can also see from panels (b) and (c) that there is a larger dispersion of wages at all levels of foreign equity participation for non-production workers compared to production workers.

If foreign investors acquire higher fractions of equity at domestic plants that pay higher wages to start with, then this sort of a selection mechanism could drive the relationship in Figure 4. To guard against such selection, Figure 5 presents the Lowess estimates that control for plant level effects. I plot average wages against the deviations from the within-plant mean value of foreign equity participation.[14] Panel (a) shows that higher levels of foreign equity participation are associated with higher average plant wages, even when the multinational status of a plant is unchanged. This means that it is not simply being foreign that brings a premium with it, but also that the size of this premium increases with the level of foreign equity paticipation. Similar to the finding in Figure 4, the Lowess estimates with fixed effects are roughly in line with the linear fit. Panels (b) and (c) show that the monotonic and positive relationship holds likewise for average production and non-production worker wages.

It is possible that the observed relationship between average wages and foreign equity participation is driven by some omitted factors. For instance, Aitken and Harrison (1999) find that foreign equity participation is positively correlated with plant productivity as measured by (log) output. If foreign plants pay their workers competitively, then this positive correlation should also be reflected in average wages. Figure 6 shows the Lowess estimates of $m(x_i)$ in equation (15) using the difference based semiparametric estimator of Yatchew (1997), which control for additional plant characteristics such as (log) value added per worker. The coefficient estimates from the difference-based semiparametric regression of (15) for the three different groups of workers are reported in Table A5 in the Appendix, along with the significance test of the nonparametric variable under the V-stat.

Panels (a) and (c) of Figure 6 confirm the earlier findings that average plant wages and average wages for non-production workers increase monotonically with foreign equity participation. The significance tests reported in Table A5 indicate that foreign equity participation is highly significant for average plant and average non-production

[14]Notice that most of the deviations from within-plant mean equity participation are positive and away from zero. This means that not only do levels of foreign ownership change at a plant over the sample period, but also that most of these changes constitute increases in the foreign ownership level. This highlights the importance of using uncensored versions of the foreign ownership variable, as the information from changes to the level of foreign ownership within the firm is lost when censored variables are used.

worker wages, with both tests delivering a p-value of zero (V-stats are 21.064 and 9.402, respectively). While the significance test for average production worker wages also returns a p-value of zero (V-stat is 5.850), panel (b) of Figure 6 casts doubt on a linear relationship for this group of workers. The estimated Lowess plot line in panel (b) is fairly flat and shows only a slight upward trend at the high end of the foreign ownership distribution. Compared to the nonparametric estimates of Figures 4 and 5, this implies that any linear relationship between wages of production workers and the level of foreign ownership is driven by other plant characteristics, such as productivity or skill composition. These results suggest that the level of foreign ownership significantly impacts the wage premium for non-production workers but it has minimal impact for production workers. Accordingly, the monotonic relationship between average plant wages and level of foreign ownership is likely to be driven by the wage premium for non-production workers only.

5.3 Estimates with Uncensored Regressors

My earlier results suggest that the size of the wage premium is affected by the level of foreign ownership. This subsection presents accurate estimates of the wage premium from the model in (16) with uncensored regressors, which not only controls for plant level effects, but also accommodates the endogeneity of foreign ownership and control variables.

My preferred set of results from system GMM estimation are reported in Table 3, where I treat all right hand side variables as potentially endogenous. This specification generates GMM style instruments for all right hand side variables, which results in close to two hundred instruments in some cases. While a larger number of instruments tends to increase efficiency, using deeper lags as instruments may weaken the strength of the instruments. In addition, instrument proliferation undermines the Hansen test, which is typically used to check instrument validity. Estimates using all available lags for the instrument set are given in columns (1), (3), and (5), while estimates using the restricted subset of instruments to the two most immediate lags are given in columns (2), (4), and (6).[15]

[15] Consistent estimation of equation (16) relies on the assumption that the idiosyncratic errors are serially uncorrelated. Test statistics for this assumption are given in Table 3 as *m1*, *m2*, and *m3* in terms of their p-values, which are tests proposed by Arellano and Bond (1991) to detect first-order, second-order, and third-order serial correlation in the *differenced* equation. Since the Arellano-Bond test statistics in Table 3 reveal second-order serial correlation, I restrict the instrument set to lags three and deeper.

The main result that comes out of Table 3 is that there exists a positive and significant relationship between foreign equity participation and average wages only for non-production workers. Column (5) indicates that a 10 percent increase in foreign equity participation leads to a 4 percent increase in the average non-production worker wage. Restricting the instrument set to lags three and four in column (6) yields an estimate of 5 percent. In the case where a plant goes from domestic ownership to being completely foreign owned (i.e. FEP goes from 0 to 100 percent), columns (5) and (6) predict the wage premium for non-production workers to be between 39 and 54 percent. Once we take into account the endogeneity of the foreign ownership variable, there is no longer an average plant wage premium due to the level of foreign ownership. This result contrasts with the FE estimates from Tables 1 and 2, which return a positive and significant foreign ownership premium for the average plant wage. Hence, simply controlling for unobserved heterogeneity at the plant level and failing to take into account other sources of endogeneity may generate considerably different results.[16]

My estimates reported in Table 3 confirm earlier findings that non-production workers are the primary beneficiaries of foreign ownership. Unlike previous studies, however, I do not find a significant wage premium for production workers, as seen from columns (3) and (4). In addition, my estimates for the hypothetical case for a plant being completely foreign owned yield larger estimates compared to earlier findings and they provide an upper bound on the estimated premium. This is due to the continuous nature of my foreign investment variable. For example, column (5) suggests that a domestic plant at which a foreign investor owns 20 percent of the equity will see the average non-production worker wage to be only 8 percent higher. However, if the foreign investor owns 80 percent of the equity, then the estimated wage premium is 32 percent. While the plant would be classified as multinational under both cases, there is a significant difference between the wage premia depending on how much of the plant equity is foreign owned. As can be seen from Figure 1, most foreign plants in Turkey have a partial degree of foreign control; the wage premia across these plants will therefore be uneven. As a result, previous studies most likely capture some estimate that lies in the range reported here and thus hide the heterogeneity in the wage premium that arises due to different levels of control.

The coefficient estimates for the controls in Table 3 are as expected, except for

[16]Note that the point I make here is not due to censoring, but due to endogeneity only. A fixed effects regression with the uncensored foreign ownership variable, not reported here, returns a significant estimate, while controlling for endogeneity via system GMM removes this significance.

(log) plant size, which seem susceptible to the specification of the instrument set. The composition of the instrument set also affects the test statistics I use to check instrument validity. The Hansen test statistics in Table 3 cannot reject the null hypothesis that the set of GMM instruments used in estimation is valid, although a large number of instruments tends to reduce the power of this test. I therefore report additionally the Sargan test of overidentifying restrictions and the Arellano and Bond (1991) test statistics for serial correlation. Both of these additional tests suggest that the sets of instruments used in the regressions are valid, although the Sargan test rejects their validity in columns (4) and (6) at the one percent confidence level.[17]

Higher levels of foreign ownership may lead to higher wage premia if plants with majority foreign control are inherently different than plants with minority control. This could arise, for instance, if a majority foreign equity participation is required for bringing in technologies from the parent firm (Lipsey and Sjoholm, 2006). In addition, Arnold and Javorcik (2009) suggest that foreign owners may substitute expatriate staff for local managers and introduce pay scales linked to performance. Gaining majority control at a plant is likely to lead to such reshuffling of the plant's labor force, especially at the administrative level, and possibly more on the job training. In order to test for such "sheepskin effects," I estimate equation (16) with an additional control, which is a dummy variable indicating majority ownership. The results from this exercise are reported in Table 4, where all right hand side variables are assumed to be endogenous.

The main result that there exists a significant wage premium only for non-production workers is confirmed by Table 4. Having majority control is far from being statistically significant in all my regressions. Column (5) indicates that conditional on having majority control, a 10 percent increase in foreign equity participation is associated with a 7 percent increase in the average wage of a non-production worker. A plant that is completely owned by foreign investors is predicted to have a wage premium of 39 percent, which matches the estimate from the same column of Table 3. Column (6) in Table 4 predicts the same wage premium to be 53 percent. The test statistics for instrument validity cannot reject the null hypothesis of exogenous instruments at the five percent confidence level, except for the Sargan statistic in columns (4) and (6). These results

[17] Roodman (2006) argues that the Sargan and Hansen tests should not be relied upon too faithfully as they are prone to weakness. While the Sargan test is not vulnerable to instrument proliferation as the Hansen test, it requires homoskedastic errors for consistency, which is rarely the case in plant level studies. Arellano and Bond (1991) also report greater power for their own proposed tests in identifying whether serial correlation renders lagged instruments invalid when compared to the Sargan and Hansen tests.

show that the positive relationship between average wages of non-production workers and the level of foreign ownership is not driven by plants under majority foreign control.

Tables 3 and 4 document that average wages of production workers are unaffected by the level of foreign control. In order to check whether such a relationship is truly nonexistent, I estimate the model in (16) with a different dependent variable. Instead of using the average yearly wage, I use the average hourly wage for production workers to calculate the wage premium. Using hourly wage data has the advantage of controlling for overtime work and can better capture the competitive wage.[18] Additionally, one reason I don't find a significant wage premium for production workers might be if foreign plants employ a greater fraction of their production workers on temporary contracts. Table A6 in the Appendix shows the results of this exercise.[19] Consistent with my earlier estimates, I find no significant premium for production workers in all of the specifications. However, I should note that both Sargan and Hansen test statistics strongly reject the validity of the instruments, which casts doubt on the reliability of these estimates.

The findings that only non-production workers benefit from multinational activity and that the wage premium depends on the level of foreign ownership can help identify which of the channels previously mentioned in the literature are at work. Arnold and Javorcik (2009) argue that while foreign owners do not alter the skill composition of labor at acquired plants, they are able to attract more experienced and motivated workers. My results suggest that multinationals attract such workers only for white collar jobs and that higher foreign equity participation is likely to reshuffle the labor force engaged in administrative work. Moreover, I interpret my findings as providing evidence for profit-sharing arguments at multinational plants. According to this branch of the literature, multinationals can afford to pay higher wages to its workers if foreign ownership is associated with higher productivity and profitability.[20] Aitken and Harrison (1999) and Takii and Ramstetter (2005) provide some evidence for the positive relationship between foreign equity participation and productivity, which seems to be the driver behind the wage premia observed at multinational plants. However, my re-

[18] One reason we are observing higher wages at foreign plants might be that workers at foreign plants might be working longer hours on a given workday or might be taking leave on a less frequent basis than their counterparts at domestic plants.

[19] The dynamic specification for this wage series seems to be clear of first-order serial correlation as suggested by the Arellano-Bond test statistic *m2*. Therefore, I also present results from regressions that use second lags and deeper as their set of instruments.

[20] See, for instance, Egger and Kreickemeier (2010).

sults suggest that profit sharing within a multinational is limited to non-production workers.

Although some existing studies consider the impact of the level of foreign ownership on the wage premium, they do not find conclusive evidence. Using panel data from Indonesian manufacturing, Lipsey and Sjoholm (2006) find that while both majority- and minority-owned foreign plants pay higher wages than domestic plants, majority-owned plants pay higher wages for white-collar workers but lower wages for blue-collar workers. However, the authors argue that none of the differences between the foreign majority and minority wages are significant at the 5% level. A similar result is reported by Aitken et al. (1996), who find, using data from Venezuela, that skilled workers receive around 4 percent higher wages at majority-owned plants compared to minority-owned plants. Hence, the current study is the first in the literature to identify systematic heterogeneity in the wage premium due to different levels of foreign ownership.

Robustness Checks

System GMM estimates are usually sensitive to the assumptions made about the variables of interest and other controls with regard to their exogeneity. These assumptions determine how the right hand side variables enter the instrument matrix in the construction of the GMM estimator and thus directly affect the number of instruments created. I previously assumed that the control variables in (16), such as skill intensity and (log) value added per worker, are potentially endogenous. This results in a large number of instruments, which can lead to an overfitting of the variables of interest. In my first round of robustness checks, I therefore provide estimates for the model where additional controls are treated as exogenous, which greatly reduces the number of instruments used in estimation.[21] The results from the baseline model in (16) with exogenous controls are reported in Table 5. Note that foreign equity participation is still assumed to be endogenous.

Table 5 confirms the main findings from the previous section, but point estimates for some variables of interest differ from those in Table 3. I again find that foreign equity participation significantly affects the average wages of non-production workers only,

[21]Strict exogeneity rules out any feedback from current or past shocks to current values of the variable, which is often not a natural restriction in the context of economic models relating to several jointly determined outcomes (Bond, 2002). While one can imagine a case where the level of foreign ownership and the skill intensity of the employees at a plant are determined concurrently, it is not as straightforward to assume that the former variable will be determined at the same time as, for instance, plant size or inputs.

but with a higher premium. Columns (5) and (6) indicate that a 10 percent increase in foreign ownership leads to an increase in the average non-production worker wage by 5.4 and 6.9 percent, respectively. The results for the controls are generally similar to the ones in Table 3, except for plant size. Under the assumption of exogeneity, plant size is negatively associated with the average plant wage (columns (1) and (2)), yet it generates a positive and significant wage premium for non-production workers (columns (5) and (6)). The Arellano-Bond test statistics confirm the presence of second-order serial correlation, justifying the use of third lags and deeper for the instrument set. However, the Sargan and Hansen statistics for overidentifying restrictions point to weaker instrument validity. This is despite the finding that the coefficients for the lagged wage term in Table 5 are typically high, which corroborates the use of a system GMM estimator as opposed to the simpler difference estimator.

I repeat the same robustness exercise, this time including a dummy variable indicating majority control.[22] The results, reported in Table 6, confirm my earlier findings. Columns (5) and (6) predict that the wage premium at a plant with 100 percent foreign equity participation is 53 and 66 percent, respectively, which are much higher estimates compared to the results in Table 4. The Hansen and Sargan test statistics in Table 6 point to weaker instrument validity, although the Arellano-Bond test statistics validate the use of third lags and deeper. As a result, treating right hand side controls in the dynamic wage model as exogenous overestimates the wage premium and undermines instrument validity. This is also suggested by Table A7, which shows the results for the average hourly wage for production workers when controls are assumed to be exogenous. The estimates for FEP and $LogWage_{t-1}$ are highly susceptible to instrument specification for this wage series, as demonstrated in the results across columns (1)-(4). Column (2) finds a marginally significant and positive effect of the level of foreign ownership on average hourly wages. However, both Sargan and Hansen test statistics strongly reject the validity of the instruments used in the estimation.

In a second round of robustness checks, I repeat all of the system GMM estimations reported here using a one-step estimator, which is not subject to the critique of downward biased standard errors in small samples.[23] The one-step results confirm my earlier findings and provide similar estimates for the wage premium for non-production workers. The estimated coefficients on FEP are almost identical for non-production workers regardless of whether the instrument set uses all available lags or is restricted

[22]Differently from the other controls, however, majority control is assumed to be endogenous.

[23]These results are available upon request in an additional appendix.

to the two most recent lags, and they suggest a wage premium of 5 percent for a 10 percent increase in foreign equity participation. Interestingly, the one-step results find a negative and significant estimate for majority control when the indicator variable is estimated along with the baseline dynamic model.[24]

6 Conclusion

A large empirical literature has identified a persistent and significant difference between average wages at multinational firms compared to their domestic counterparts. There exists a wage premium to being multinational even after controlling for selection effects whereby foreign investors cherry-pick the plants they acquire. At the time of acquisition or subsequently, the individual characteristics and experiences of foreign investors are likely to impact the degree of control they want to exercise. The level of control that foreign owners choose at newly acquired plants may also have differential effects for its production and non-production workers. This requires that empirical studies that explore the relationship between wages and foreign ownership would be better equipped if they explicitly consider different levels of foreign equity participation and account for the endogeneity of foreign ownership. Most of the previous literature worked with binary variables to indicate foreign ownership, which might mislead researchers' understanding of the impact of foreign ownership on wages. Estimation with censored variables on the right hand side returns biased results even after controlling for individual level effects. One implication is that one cannot readily compare estimates from country studies with each other, as the distribution of foreign ownership shares across firms and thresholds used in the definitions of foreign ownership are likely to vary across countries.

My results provide more accurate estimates for the effect of foreign ownership on wages by using continuous data for the variable under study at the same time as controlling for endogeneity. This allows me to identify the heterogeneity in the wage premium that arises due to different levels of foreign equity participation. I estimate that a 10% increase in the level of foreign ownership is associated with about a 4% increase in the average wage of a non-production worker. My results suggest that the identified wage

[24] As a further test of whether my results are driven by the variation in foreign ownership at a certain subset of foreign plants, I experimented with system GMM regressions of the dynamic wage model in (16) using only the subsamples of plants under minority and majority control separately. However, this exercise runs into the problem of cutting the sample of plants used in estimation by around a half. As a result, the number of plants (groups) used in the GMM estimation gets closer to the number of instruments generated, which severely distorts inference. In my estimations on each subsample, I frequently observe a Hansen test statistic of 1.0, which indicates the severity of this problem.

premium at multinationals is primarily driven by higher pay for the non-production workers. I do not find a wage premium for production workers across a variety of empirical settings. In addition, failing to address the endogeneity of foreign ownership returns misestimates of the wage premium. A more informed choice of explanatory variables and econometric specification are thus crucial to better understand both the impact and the size of the foreign ownership wage premium. The heterogeneity identified in this paper also raises several issues for further research, especially theoretical models of foreign direct investment. Why are higher levels of foreign equity participation associated with higher wages? In addition, why is such a relationship only observed for non-production workers? Future research that investigates these questions would be welcome.

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Figure 1: Censoring the Level of Foreign Ownership

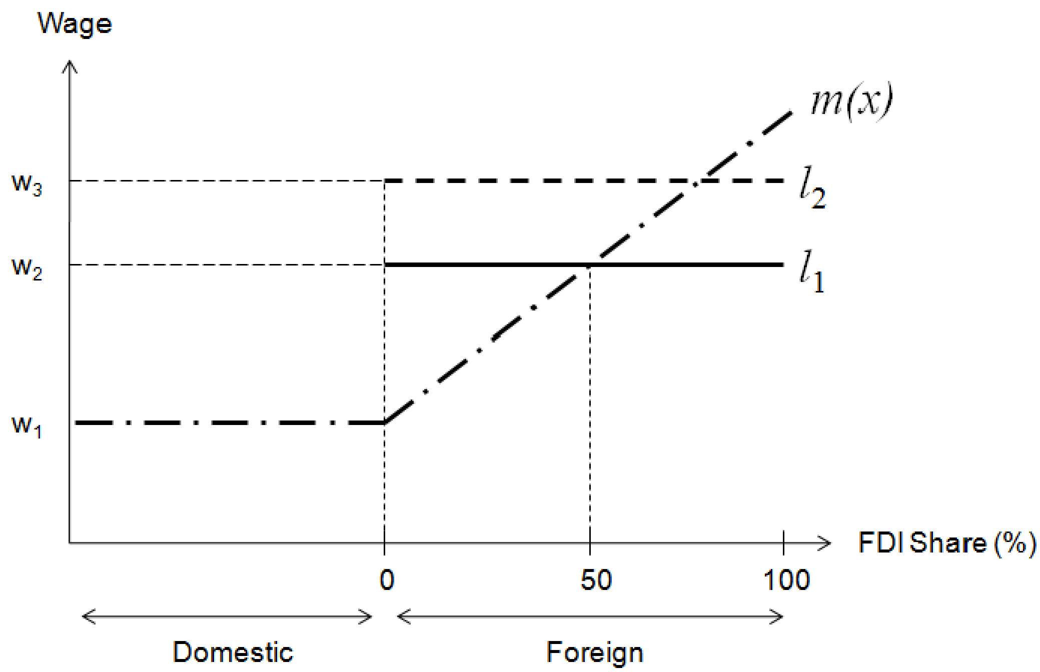


Figure 2: Distribution of Foreign Ownership Shares at the Plant Level, 1993-2001

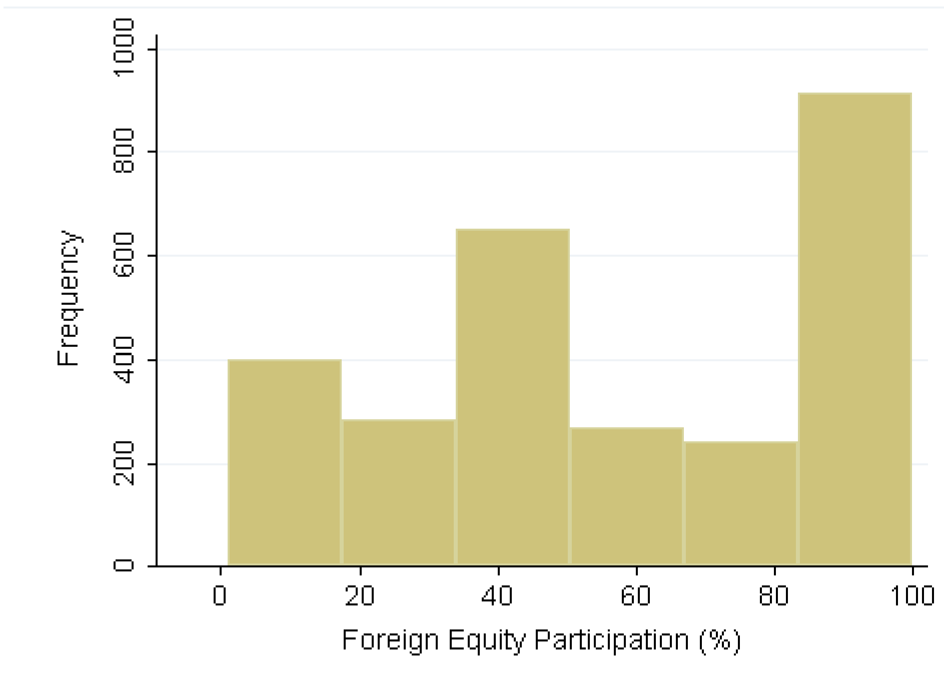
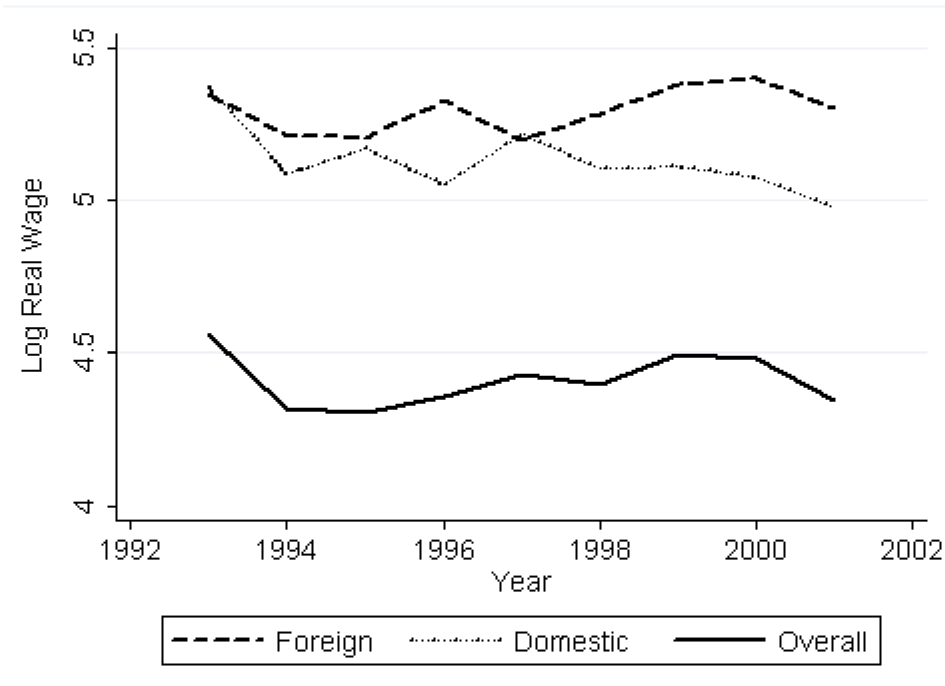


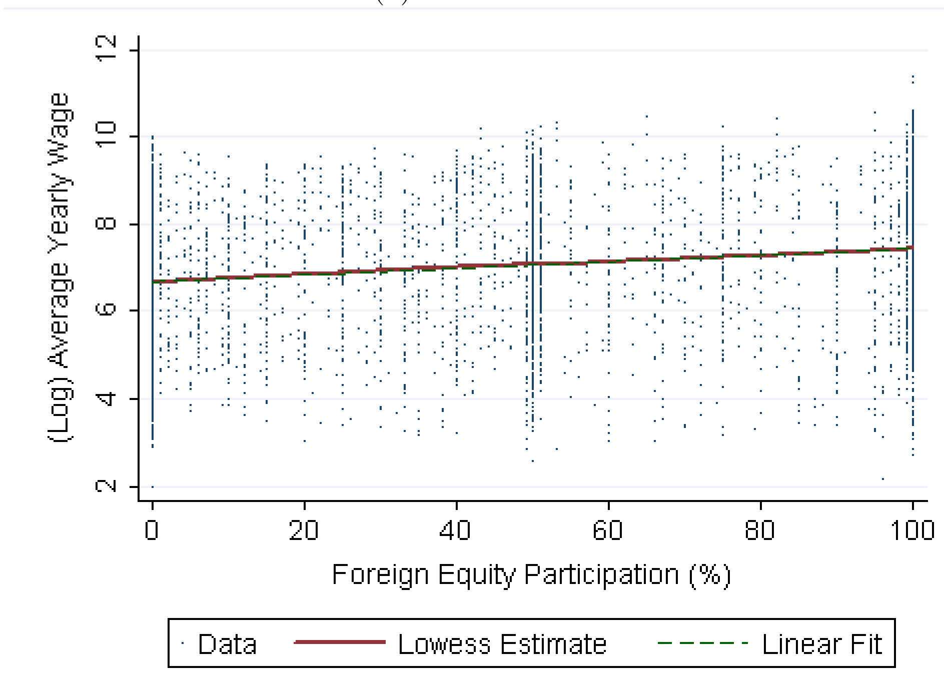
Figure 3: Comparison of Wages Across Plants that Experienced a Takeover



Notes: "Foreign" and "Domestic" refer to plants that were subject to a foreign takeover at one point in the sample period and depict wages at these plants when they were under foreign and domestic control, respectively. "Overall" depicts the pattern from the pooled sample of all plants.

Figure 4: Nonparametric Estimates of the Relationship between Average Yearly Wage and Share of Foreign Ownership: Pooled OLS Regression

(a) All Workers



(b) Production Workers

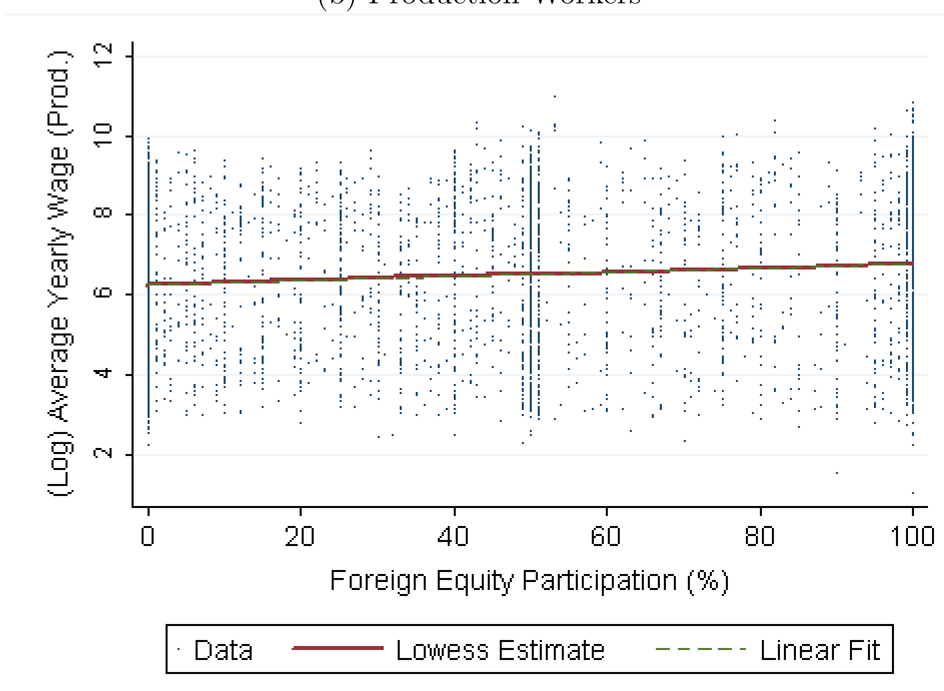


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Figure 4 (continued): Nonparametric Estimates of the Relationship between Average Yearly Wage and Share of Foreign Ownership: Pooled OLS Regression

(c) Non-production Workers

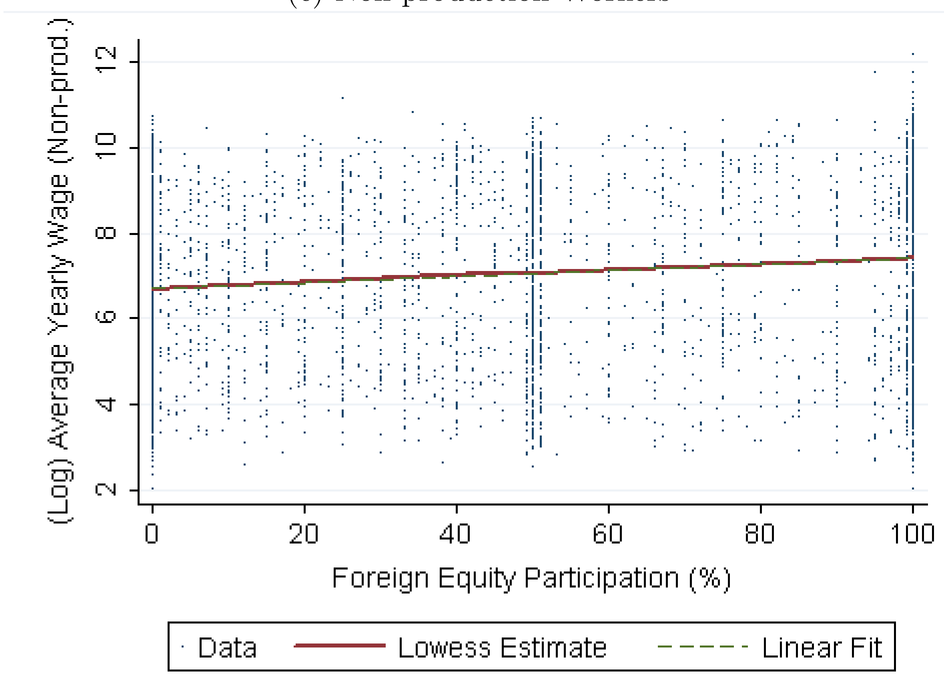
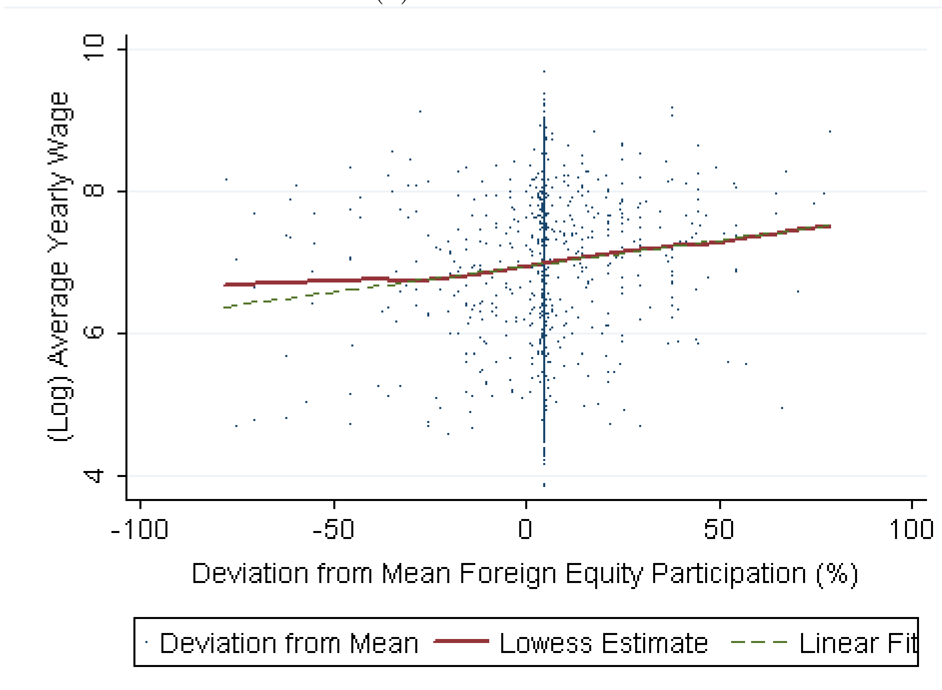


Figure 5: Nonparametric Estimates of the Relationship between Average Yearly Wage and Share of Foreign Ownership: Fixed Effects Regression

(a) All Workers



(b) Production Workers

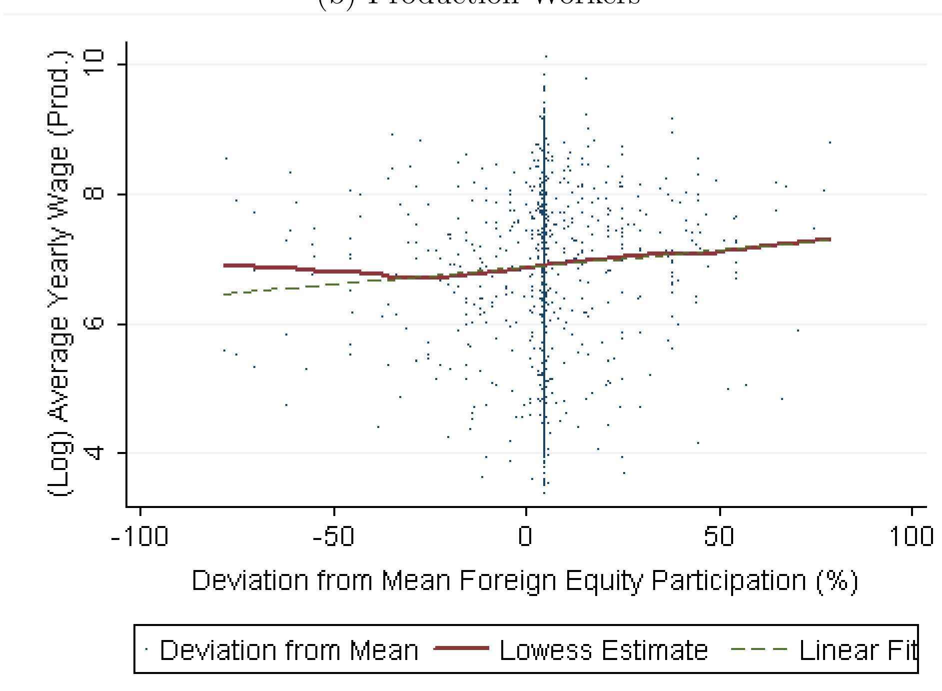


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Figure 5 (continued): Nonparametric Estimates of the Relationship between Average Yearly Wage and Share of Foreign Ownership: Fixed Effects Regression

(c) Non-production Workers

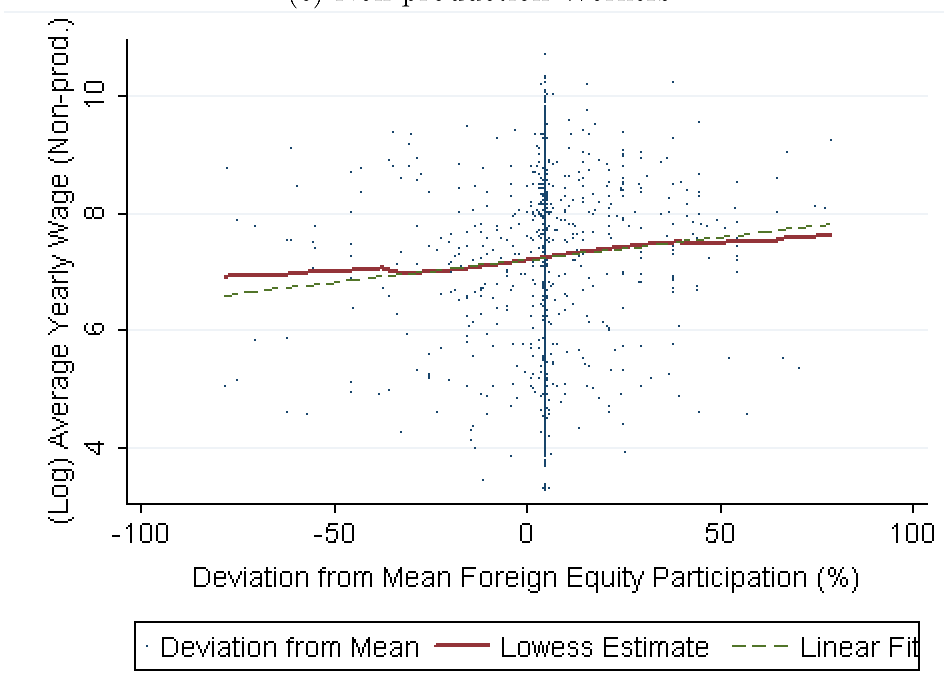
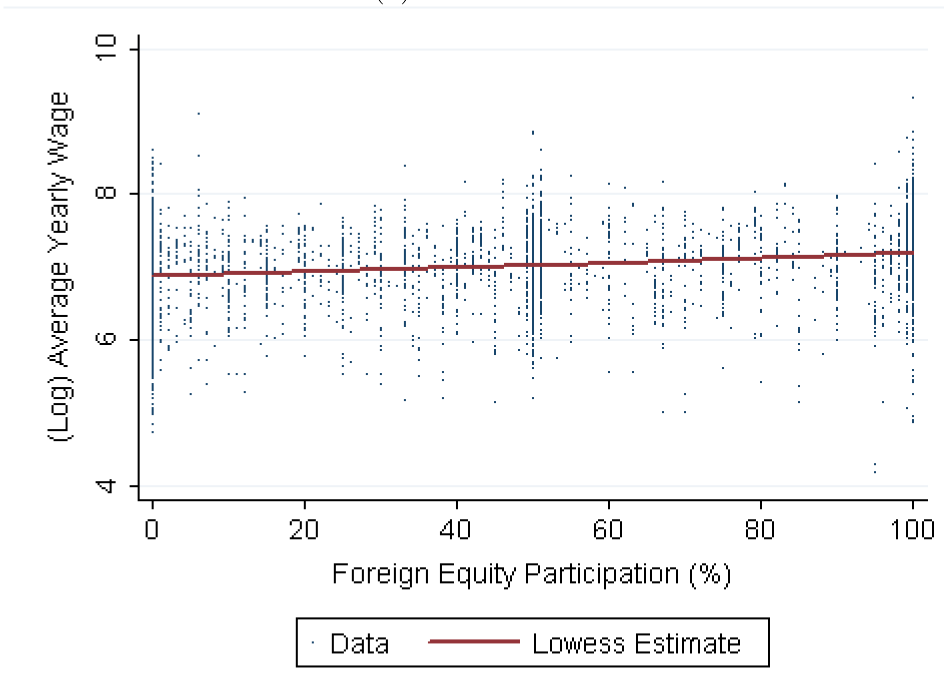


Figure 6: Semiparametric Estimates of the Relationship between Average Yearly Wage and Share of Foreign Ownership: Fixed Effects Regression

(a) All Workers



(b) Production Workers

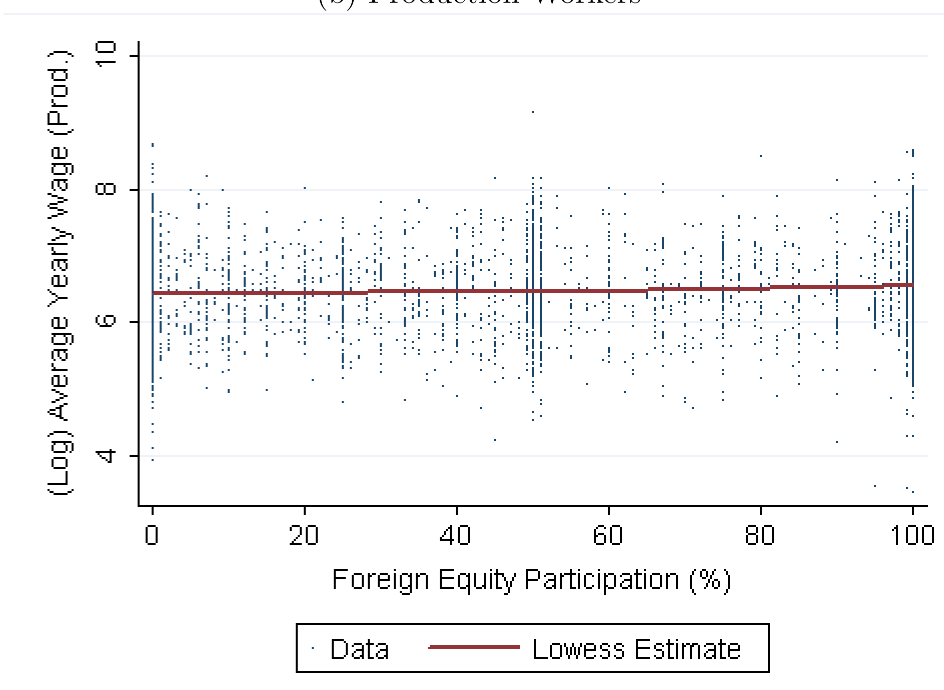


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Figure 6 (continued): Semiparametric Estimates of the Relationship between Average Yearly Wage and Share of Foreign Ownership: Fixed Effects Regression

(c) Non-production Workers

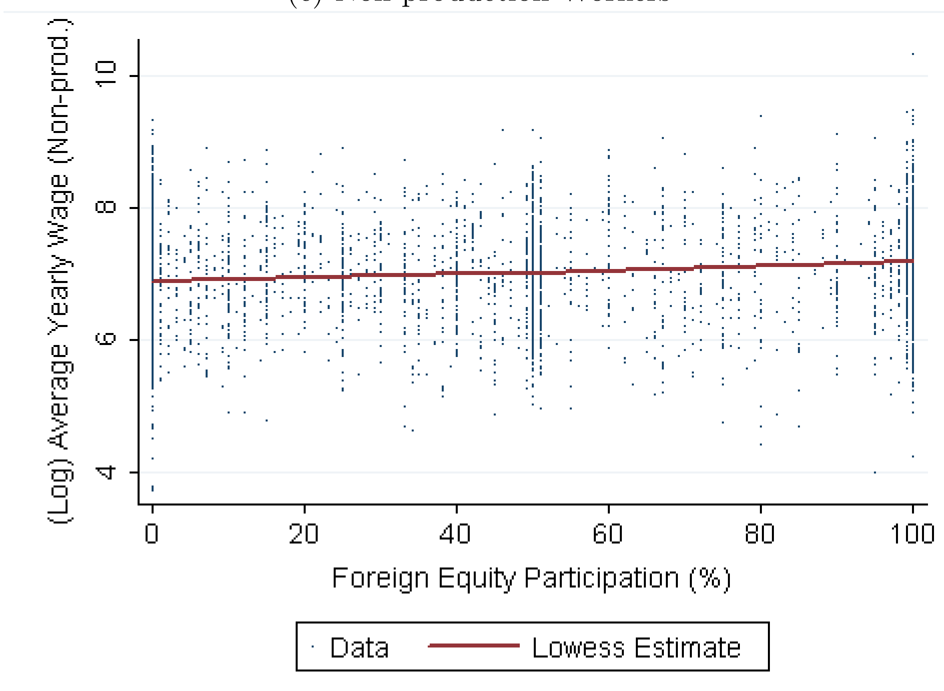

Table 1: OLS and FE Results: Wages and Multinational Status Defined at Various Thresholds
Dependent Variable: Log Average Yearly Wage

		OLS Estimates			FE Estimates		
		All Workers	Production Workers	Non-production Workers	All Workers	Production Workers	Non-production Workers
Foreign Equity Participation Threshold		(1)	(2)	(3)	(4)	(5)	(6)
0%	(a)	.4839 (.0212)***	.1708 (.0164)***	.3728 (.0217)***	.1121 (.0214)***	.0441 (.0270)	.1802 (.0409)***
15%	(b)	.5064 (.0228)***	.1839 (.0176)***	.4046 (.0233)***	.1144 (.0224)***	.0468 (.0280)*	.2142 (.0427)***
30%	(c)	.5246 (.0242)***	.1946 (.0188)***	.4133 (.0251)***	.1165 (.0241)***	.0376 (.0288)	.2001 (.0461)***
50%	(d)	.5677 (.0293)***	.2167 (.0235)***	.4513 (.0311)***	.1354 (.0307)***	.0611 (.0398)	.2505 (.0595)***

Notes: This table reports the coefficient estimates for the censored foreign ownership variable in the model in (12). The full set of results for the OLS and FE regressions are in the Appendix, Tables A3 and A4, respectively. All standard errors are corrected for heteroskedasticity, clustered at the plant level. Coefficients are given in the first line; standard errors in parentheses; *, **, *** indicate significance at the 10%, 5% and 1% level, respectively. All regressions include (log) plant size, skill intensity, ratio of production workers, (log) value added per worker, and (log) electricity as controls. OLS regressions include sector, region, and year dummies, and FE regressions include individual plant effects and year dummies as additional controls.

Table 2: OLS and FE Results: Wages and Multinational Status Defined at Various Intervals
Dependent Variable: Log Average Yearly Wage

	OLS Estimates			FE Estimates		
	All Workers	Production Workers	Non-production Workers	All Workers	Production Workers	Non-production Workers
	(1)	(2)	(3)	(4)	(5)	(6)
FDI dummy	.3231	.0864	.3033	.0775	.0625	.2008
(interval 15-30%)	(.0469)***	(.0376)**	(.0514)***	(.0342)**	(.0598)	(.0998)**
FDI dummy	.4191	.1440	.3220	.0850	.0131	.1351
(interval 31-50%)	(.0377)***	(.0294)***	(.0399)***	(.0268)***	(.0417)	(.0619)**
FDI dummy	.5846	.2221	.4656	.1530	.0662	.2817
(interval 51-100%)	(.0294)***	(.0236)***	(.0311)***	(.0320)***	(.0396)*	(.0603)***
Log Plant Size	.1977	.0628	.0962	.0138	-.0246	.0141
	(.0038)***	(.0042)***	(.0059)***	(.0056)**	(.0067)***	(.0099)
Skill Intensity	.0040	.0067	.0037	.0003	.0051	.0017
	(.0001)***	(.0002)***	(.0002)***	(.0001)***	(.0002)***	(.0002)***
Ratio of Production	.0008	.0083	.0116	-.0006	.0070	.0102
Workers	(.0003)**	(.0024)***	(.0037)***	(.0005)	(.002)***	(.0035)***
Log Value Added	.2013	.0673	.0756	.0897	.0364	.0241
per Worker	(.0037)***	(.0028)***	(.0035)***	(.0025)***	(.0029)***	(.0040)***
Log Electricity	.0093	.1091	.1363	.0043	.1223	.1524
	(.0013)***	(.0026)***	(.0041)***	(.0009)***	(.0026)***	(.0041)***
Model Effects	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.9151	0.8956	0.8416	0.8684	0.8847	0.8292
N	91,555	91,392	80,975	91,555	91,392	80,975

Notes: All standard errors are corrected for heteroskedasticity, clustered at the plant level. Coefficients are given in the first line; standard errors in parentheses; *, **, *** indicate significance at the 10%, 5% and 1% level, respectively. Model effects include sector, region, and year dummies for the OLS regressions, and individual plant effects and year dummies for the FE regressions. All regressions include a constant term. Reference category: FDI dummy (interval 0-15%).

Table 3: Two-Step System GMM Results: Wages and Foreign Ownership (Endogenous Controls)
Dependent Variable: Log Average Yearly Wage

	All Workers		Production Workers		Non-Production Workers	
	(1)	(2)	(3)	(4)	(5)	(6)
Log Wage $_{t-1}$	.6898	.7395	.3378	.3303	.3880	.4152
	(.0554)***	(.0630)***	(.0521)***	(.0646)***	(.0592)***	(.0749)***
Foreign Equity	.0005	-.00008	.0013	.0007	.0039	.0054
Participation (%)	(.0006)	(.0008)	(.0008)	(.0011)	(.0011)***	(.0013)***
Log Plant Size	.0078	.0445	.0477	.1432	-.0341	-.0333
	(.0277)	(.0338)	(.0528)	(.0649)**	(.0688)	(.0832)
Skill Intensity	.0031	.0039	.0082	.0084	.0040	.0047
	(.0009)***	(.0011)***	(.0019)***	(.0025)***	(.0023)*	(.0026)*
Ratio of Production	.0028	-.0015	.0985	.1025	.0763	.0753
Workers	(.0085)	(.0097)	(.0389)**	(.0475)**	(.0232)***	(.0198)***
Log Value Added per	.1048	.0928	.1788	.2122	.1091	.0247
Worker	(.0334)***	(.0480)*	(.0507)***	(.0745)***	(.0765)	(.0957)
Log Input	.0289	-.0037	.1065	.0431	.1668	.1893
	(.0287)	(.0372)	(.0450)**	(.0567)	(.0692)**	(.0870)**
m1 (Pr>z)	0.000	0.000	0.000	0.000	0.000	0.000
m2 (Pr>z)	0.017	0.018	0.117	0.144	0.015	0.017
m3 (Pr>z)	0.736	0.791	0.219	0.209	0.801	0.902
Sargan	0.036	0.032	0.122	0.001	0.109	0.005
Hansen	0.424	0.610	0.314	0.146	0.141	0.154
Number of Instruments	197	127	197	127	197	127
Instrument Set	lags 3+	lags 3 and 4	lags 3+	lags 3 and 4	lags 3+	lags 3 and 4
N	3513	3513	3484	3484	3233	3233

Notes: Year dummies and a constant term included in all models. Controls treated as *endogenous*. Robust standard errors in parentheses, clustered at the plant level and adjusted for Windmeijer's correction; *, **, *** indicate significance at the 10%, 5% and 1% level, respectively. m1, m2, and m3 are Arellano-Bond tests for first-order, second-order, and third-order serial correlation, asymptotically $N(0,1)$. Sargan and Hansen are tests of the overidentifying restrictions for the GMM estimators, asymptotically χ^2; p-value is reported. These tests use the minimized value of the corresponding two-step GMM estimators.

Table 4: Two-Step System GMM Results: Wages and Foreign Majority Ownership (Endogenous Controls)
Dependent Variable: Log Average Yearly Wage

	All Workers		Production Workers		Non-Production Workers	
	(1)	(2)	(3)	(4)	(5)	(6)
Log Wage $_{t-1}$	.6675	.7280	.3375	.3255	.3667	.4265
	(.0525)***	(.0627)***	(.0537)***	(.0646)***	(.0588)***	(.0723)***
Foreign Equity	.0003	.0014	.0027	.0031	.0072	.0081
Participation (%)	(.0011)	(.0014)	(.0021)	(.0027)	(.0025)***	(.0030)***
Majority Share Dummy	.0218	-.1042	-.1204	-.2354	-.3285	-.2766
	(.0809)	(.1030)	(.1711)	(.2261)	(.2055)	(.2410)
Log Plant Size	.0068	.0443	.0496	.1326	-.0458	-.0465
	(.0257)	(.0327)	(.0498)	(.0620)**	(.0713)	(.0816)
Skill Intensity	.0029	.0039	.0078	.0076	.0033	.0041
	(.0009)***	(.0010)***	(.0019)***	(.0025)***	(.0023)	(.0027)
Ratio of Production	.0015	-.0028	.0970	.1019	.0768	.0760
Workers	(.0085)	(.0098)	(.0390)**	(.0490)**	(.0249)***	(.0208)***
Log Value Added per	.1187	.0889	.1780	.2299	.1185	.0285
Worker	(.0307)***	(.0421)**	(.0494)***	(.0725)***	(.0756)	(.0893)
Log Input	.0298	-.0005	.0997	.0387	.1738	.1841
	(.0256)	(.0346)	(.0421)**	(.0568)	(.0666)***	(.0807)**
m1 (Pr>z)	0.000	0.000	0.000	0.000	0.000	0.000
m2 (Pr>z)	0.017	0.019	0.127	0.170	0.020	0.013
m3 (Pr>z)	0.681	0.789	0.203	0.190	0.746	0.879
Sargan	0.090	0.127	0.065	0.000	0.156	0.006
Hansen	0.384	0.651	0.188	0.094	0.180	0.230
Number of Instruments	224	144	224	144	224	144
Instrument Set	lags 3+	lags 3 and 4	lags 3+	lags 3 and 4	lags 3+	lags 3 and 4
N	3513	3513	3484	3484	3233	3233

Notes: Year dummies and a constant term included in all models. Controls treated as *endogenous*. Robust standard errors in parentheses, clustered at the plant level and adjusted for Windmeijer's correction. See notes to Table 3.

Table 5: Two-Step System GMM Results: Wages and Foreign Ownership (Exogenous Controls)
Dependent Variable: Log Average Yearly Wage

	All Workers		Production Workers		Non-Production Workers	
	(1)	(2)	(3)	(4)	(5)	(6)
Log Wage $_{t-1}$	.6181	.6496	.5095	.5114	.5124	.5317
	(.0988)***	(.0961)***	(.0566)***	(.0634)***	(.0755)***	(.0846)***
Foreign Equity	.0014	.0011	.0016	.0008	.0054	.0067
Participation (%)	(.0009)	(.0011)	(.0013)	(.0015)	(.0018)***	(.0020)***
Log Plant Size	-.0294	-.0280	.0249	.0274	.0466	.0467
	(.0149)**	(.0147)*	(.0186)	(.0195)	(.0242)*	(.0234)**
Skill Intensity	.0018	.0017	.0085	.0087	.0032	.0029
	(.0006)***	(.0006)***	(.0011)***	(.0011)***	(.0013)**	(.0013)**
Ratio of Production	.0021	.0020	.0942	.0941	.1156	.1159
Workers	(.0049)	(.0047)	(.0225)***	(.0220)***	(.0297)***	(.0274)***
Log Value Added per	.1041	.1026	.0697	.0720	.0916	.0855
Worker	(.0231)***	(.0225)***	(.0198)***	(.0208)***	(.0243)***	(.0241)***
Log Input	.0660	.0591	.0535	.0527	.0553	.0524
	(.0163)***	(.0162)***	(.0156)***	(.0168)***	(.0200)***	(.0198)***
m1 (Pr>z)	0.000	0.000	0.000	0.000	0.000	0.000
m2 (Pr>z)	0.040	0.042	0.021	0.021	0.011	0.014
m3 (Pr>z)	0.729	0.765	0.184	0.193	0.883	0.842
Sargan	0.007	0.001	0.089	0.002	0.000	0.000
Hansen	0.070	0.029	0.011	0.000	0.006	0.004
Number of Instruments	67	47	67	47	67	47
Instrument Set	lags 3+	lags 3 and 4	lags 3+	lags 3 and 4	lags 3+	lags 3 and 4
N	3513	3513	3484	3484	3233	3233

Notes: Year dummies and a constant term included in all models. Controls treated as *exogenous*. Robust standard errors in parentheses, clustered at the plant level and adjusted for Windmeijer's correction. See notes to Table 3.

Table 6: Two-Step System GMM Results: Wages and Foreign Majority Ownership (Exogenous Controls)
Dependent Variable: Log Average Yearly Wage

	All Workers		Production Workers		Non-Production Workers	
	(1)	(2)	(3)	(4)	(5)	(6)
Log Wage $_{t-1}$	.5892	.6574	.4669	.4815	.4510	.5045
	(.0840)***	(.0873)***	(.0599)***	(.0631)***	(.0768)***	(.0821)***
Foreign Equity	.0013	.0018	.0044	.0042	.0087	.0106
Participation (%)	(.0014)	(.0016)	(.0027)	(.0030)	(.0032)***	(.0034)***
Majority Share Dummy	.0171	-.0523	-.1861	-.2810	-.3358	-.4008
	(.1011)	(.1269)	(.2145)	(.2412)	(.2556)	(.2562)
Log Plant Size	-.0248	-.0267	.0239	.0269	.0454	.0393
	(.0151)	(.0147)*	(.0194)	(.0199)	(.0239)*	(.0234)*
Skill Intensity	.0020	.0017	.00837	.0085	.0038	.0033
	(.0005)***	(.0006)***	(.0012)***	(.0011)***	(.0013)***	(.0013)***
Ratio of Production	.0023	.0020	.0941	.0957	.1216	.1173
Workers	(.0046)	(.0045)	(.0227)***	(.0233)***	(.0322)***	(.0290)***
Log Value Added per	.1097	.1018	.0703	.0812	.1050	.0922
Worker	(.0214)***	(.0213)***	(.0202)***	(.0208)***	(.0242)***	(.0241)***
Log Input	.0664	.0559	.0560	.0511	.0606	.0559
	(.0138)***	(.0144)***	(.0163)***	(.0172)***	(.0206)***	(.0202)***
m1 (Pr>z)	0.000	0.000	0.000	0.000	0.000	0.000
m2 (Pr>z)	0.037	0.037	0.039	0.034	0.019	0.014
m3 (Pr>z)	0.691	0.770	0.161	0.165	0.984	0.945
Sargan	0.091	0.025	0.079	0.001	0.002	0.000
Hansen	0.102	0.111	0.004	0.001	0.017	0.021
Number of Instruments	94	64	94	64	94	64
Instrument Set	lags 3+	lags 3 and 4	lags 3+	lags 3 and 4	lags 3+	lags 3 and 4
N	3513	3513	3484	3484	3233	3233

Notes: Year dummies and a constant term included in all models. Controls treated as *exogenous*. Robust standard errors in parentheses, clustered at the plant level and adjusted for Windmeijer's correction. See notes to Table 3.

Appendix

Table A1: Summary Statistics of the Variables Used in the Analysis by Type of Ownership

		Foreign	*Domestic*
Foreign Equity Participation (%)	*N*	3140	91434
	Mean	60.12	0
	Std. Dev.	32.39	0
Average Plant Wage (Turkish Liras)	*N*	3140	91434
	Mean	3712.06	1038.28
	Std. Dev.	5534.14	1651.31
Average Wage for Production Workers (Turkish Liras)	*N*	3120	91225
	Mean	2783.73	972.94
	Std. Dev.	4562.69	1550.13
Average Wage for Non-Production Workers (Turkish Liras)	*N*	3003	79656
	Mean	5816.05	1487.05
	Std. Dev.	9876.44	2807.87
(Log) Plant Size	*N*	3140	91434
	Mean	5.01	3.68
	Std. Dev.	1.27	1.08
Skill Intensity (%)	*N*	3126	91201
	Mean	30.78	19.86
	Std. Dev.	21.41	17.12
Ratio of Production Workers	*N*	3140	91434
	Mean	0.78	1.54
	Std. Dev.	1.58	5.91
(Log) Value Added per Worker	*N*	3102	90049
	Mean	8.69	7.24
	Std. Dev.	1.76	1.75
(Log) Input	*N*	3140	91414
	Mean	13.97	11.62
	Std. Dev.	2.34	2.33

Notes: A foreign plant is defined as a manufacturing plant which has any positive ratio of foreign equity in the plant's ownership. In the sample, the minimum share of foreign ownership was 1% and the maximum share was 100%.

Table A2: Foreign Presence in the Turkish Manufacturing Sector

Year	*Number of Foreign Plants*	*Number of Domestic Plants*	*Total Number of Plants*	*Foreign Presence (%)*	*Average Share of Foreign Ownership at Foreign Plants (%)*
1993	301	10,266	10,567	2.85	58.78
1994	312	9,815	10,127	3.08	58.95
1995	325	9,904	10,229	3.18	59.96
1996	326	10,264	10,590	3.08	58.48
1997	362	11,003	11,365	3.19	57.04
1998	416	11,905	12,321	3.38	59.25
1999	406	10,856	11,262	3.61	60.08
2000	414	10,700	11,114	3.73	62.01
2001	439	10,872	11,311	3.88	64.33

Notes: A foreign plant is defined as a manufacturing plant which has any positive ratio of foreign equity in the plant's ownership. In the sample, the minimum share of foreign ownership was 1% and the maximum share was 100%. *Foreign Presence* is the ratio of *Number of Foreign Plants* to *Total Number of Plants*.

Table A3: OLS Results: Wages and Multinational Status Defined at Various Thresholds
Dependent Variable: Log Average Yearly Wage

	All Workers				Production Workers				Non-production Workers			
FDI Threshold	(1) 0%	(2) 15%	(3) 30%	(4) 50%	(5) 0%	(6) 15%	(7) 30%	(8) 50%	(9) 0%	(10) 15%	(11) 30%	(12) 50%
FDI Dummy	.4839 (.0212)***	.5064 (.0228)***	.5246 (.0242)***	.5677 (.0293)***	.1708 (.0164)***	.1839 (.0176)***	.1946 (.0188)***	.2167 (.0235)***	.3728 (.0217)***	.4046 (.0233)***	.4133 (.0251)***	.4513 (.0311)***
Log Plant Size	.1959 (.0038)***	.1975 .(0038)***	.1989 (.0038)***	.2016 (.0038)***	.0623 (.0042)***	.0627 (.0042)***	.0631 (.0042)***	.0641 (.0042)***	.0952 (.0059)***	.0962 (.0059)***	.0974 (.0058)***	.0996 (.0058)***
Skill Intensity	.0040 (.0001)***	.0040 (.0001)***	.0040 (.0001)***	.0041 (.0001)***	.0067 (.0001)***	.0067 (.0001)***	.0067 (.0001)***	.0067 (.0002)***	.0037 (.0002)***	.0037 (.0002)***	.0037 (.0002)***	.0037 (.0002)***
Ratio of Production Workers	.0007 (.0003)**	.0008 (.0003)**	.0008 (.0003)**	.0008 (.0003)**	.0083 (.0023)***	.0083 (.0024)***	.0083 (.0024)***	.0083 (.0024)***	.0116 (.0037)***	.0116 (.0036)***	.0116 (.0037)***	.0116 (.0037)***
Log Value Added per Worker	.2011 (.0037)***	.2016 (.0037)***	.2022 (.0037)***	.2047 (.0037)***	.0674 (.0028)***	.0674 (.0028)***	.0675 (.0028)***	.0684 (.0028)***	.0758 (.0035)***	.0758 (.0035)***	.0765 (.0035)***	.0785 (.0035)***
Log Electricity	.0091 (.0013)***	.0092 (.0013)***	.0093 (.0013)***	.0093 (.0013)***	.1090 (.0026)***	.1091 (.0026)***	.1091 (.0026)***	.1091 (.0026)***	.1362 (.0041)***	.1362 (.0041)***	.1363 (.0041)***	.1363 (.0041)***
Sector/Year/ Region Dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.9151	0.9150	0.9149	0.9144	0.8955	0.8955	0.8956	0.8955	0.8415	0.8416	0.8415	0.8412
N	91555	91555	91555	91555	91392	91392	91392	91392	80975	80975	80975	80975

Notes: All standard errors are corrected for heteroskedasticity (cluster at plant level). Coefficients are given in the first line; standard errors in parentheses; *, **, *** indicate significance at the 10%, 5% and 1% level, respectively. All regressions include a constant term.

Table A4: FE Results: Wages and Multinational Status Defined at Various Thresholds
Dependent Variable: Log Average Yearly Wage

	All Workers				Production Workers				Non-production Workers			
FDI Threshold	(1) 0%	(2) 15%	(3) 30%	(4) 50%	(5) 0%	(6) 15%	(7) 30%	(8) 50%	(9) 0%	(10) 15%	(11) 30%	(12) 50%
FDI Dummy	.1121 (.0214)***	.1144 (.0224)***	.1165 (.0241)***	.1354 (.0307)***	.0441 (.0270)	.0468 (.0280)*	.0376 (.0288)	.0611 (.0398)	.1802 (.0409)***	.2142 (.0427)***	.2001 (.0461)***	.2505 (.0595)***
Log Plant Size	.0136 (.0056)**	.0136 (.0056)**	.0138 (.0056)***	.0143 (.0056)***	-.0247 (.0067)***	-.0247 (.0067)***	-.0246 (.0067)***	-.0244 (.0067)***	.0139 (.0099)	.0137 (.0099)	.0143 (.0099)	.0151 (.0099)
Skill Intensity	.0003 (.00008)***	.0003 (.00008)***	.0003 (.00008)***	.0003 (.00008)***	.0051 (.0002)***	.0051 (.0002)***	.0051 (.0002)***	.0051 (.0002)***	.0017 (.0002)***	.0017 (.0002)***	.0017 (.0002)***	.0017 (.0002)***
Ratio of Production Workers	-.0006 (.0005)	-.0006 (.0005)	-.0006 (.0005)	-.0006 (.0005)	.0070 (.0020)***	.0070 (.0020)***	.0070 (.0020)***	.0070 (.0021)***	.0102 (.0031)***	.0102 (.0035)***	.0102 (.0035)***	.0102 (.0035)***
Log Value Added per Worker	.0897 (.0025)***	.0897 (.0025)***	.0897 (.0025)***	.0897 (.0025)***	.0364 (.0029)***	.0364 (.0029)***	.0364 (.0029)***	.0364 (.0029)***	.0242 (.0040)***	.0242 (.0040)***	.0242 (.0040)***	.0241 (.0040)***
Log Electricity	.0044 (.0010)***	.0044 (.0010)***	.0044 (.0010)***	.0043 (.0010)***	.1223 (.0026)***	.1222 (.0026)***	.1222 (.0026)***	.1223 (.0026)***	.1524 (.0041)***	.1524 (.0041)***	.1524 (.0041)***	.1524 (.0041)***
Time and Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Overall R^2	0.8683	0.8681	0.8680	.8677	0.8846	0.8846	0.8845	0.8846	0.8288	0.8289	0.8287	0.8285
N	91,555	91,555	92,887	92,887	91,392	91,392	91,392	91,392	80,975	80,975	80,975	80,975

Notes: All standard errors are corrected for heteroskedasticity (cluster at plant level). Coefficients are given in the first line; standard errors in parentheses; *, **, *** indicate significance at the 10%, 5% and 1% level, respectively. All regressions include a constant term.

Table A5: Results from Difference-Based Semiparametric Regression
Dependent Variable: (Log) Average Yearly Wage

	All Workers (1)	Production Workers (2)	Non-Production Workers (3)
Log Plant Size	.0555	.0519	.0073
	(.0249)**	(.0318)	(.0386)
Skill Intensity	.0044	.0109	.0052
	(.0005)***	(.0007)***	(.0008)***
Ratio of Production Workers	.0059	.0807	.1045
	(.0053)	(.0067)***	(.0082)***
Log Value Added per Worker	.3095	.1514	.1605
	(.0124)***	(.0158)***	(.0197)***
Log Input	.0959	.0659	.0950
	(.0108)***	(.0138)***	(.0172)***
Year Dummies	Yes	Yes	Yes
V-stat	21.064	5.850	9.402
(p-value)	(0.000)	(0.000)	(0.000)
R^2	0.8941	0.8722	0.8431
N	4237	4217	4042

Notes: *, **, *** indicate significance at the 10%, 5% and 1% level, respectively. V-stat is a significance test of the nonparametric component in the regression, foreign equity participation, and is asymptotically $N(0, 1)$. See Yatchew (1997). Both the test statistic and corresponding p-value are reported.

Table A6: Two-Step System GMM Results: Hourly Wages and Foreign Ownership (Endogenous Controls)
Dependent Variable: Log Average Hourly Wage for Production Workers

	(1)	(2)	(3)	(4)
Log Wage $_{t-1}$	.1246	.0971	.2665	.2516
	(.0195)***	(.0193)***	(.0363)***	(.0410)***
Foreign Equity	-.0007	.0001	.0014	.0005
Participation (%)	(.0015)	(.0018)	(.0014)	(.0018)
Log Plant Size	.0051	.0497	-.0233	.0325
	(.0740)	(.0940)	(.0843)	(.0969)
Skill Intensity	.0096	.0120	.0080	.0093
	(.0026)***	(.0029)***	(.0030)***	(.0039)**
Ratio of Production	.3814	.4183	.4528	.5150
Workers	(.1676)**	(.2027)**	(.1757)***	(.1986)***
Log Value Added per	.2025	.1813	.0287	.0133
Worker	(.0800)**	(.0896)**	(.0772)	(.1242)
Log Input	.0995	.0909	.1749	.1634
	(.0644)	(.0805)	(.0665)***	(.0836)*
m1 (Pr>z)	0.000	0.000	0.000	0.000
m2 (Pr>z)	0.573	0.353	0.957	0.742
m3 (Pr>z)	0.194	0.216	0.107	0.089
Sargan	0.000	0.000	0.000	0.000
Hansen	0.000	0.000	0.001	0.000
Number of Instruments	253	148	197	127
Instrument Set	lags 2+	lags 2 and 3	lags 3+	lags 3 and 4
N	3474	3474	3474	3474

Notes: Year dummies and a constant term included in all models. Controls treated as *endogenous*. Robust standard errors in parentheses, clustered at the plant level and adjusted for Windmeijer's correction. See notes to Table 3.

Table A7: Two-Step System GMM Results: Hourly Wages and Foreign Ownership (Exogenous Controls)
Dependent Variable: Log Average Hourly Wage for Production Workers

	(1)	(2)	(3)	(4)
Log Wage $_{t-1}$	.1026	.0738	.5489	.5610
	(.0228)***	(.0221)***	(.0474)***	(.0490)***
Foreign Equity	.0028	.0038	.0006	.0005
Participation (%)	(.0020)	(.0022)*	(.0021)	(.0025)
Log Plant Size	-.1763	-.1841	-.0238	-.0271
	(.0466)***	(.0472)***	(.0373)	(.0369)
Skill Intensity	.0080	.0080	.0081	.0078
	(.0022)***	(.0022)***	(.0022)***	(.0023)***
Ratio of Production	.4380	.4601	.4290	.4337
Workers	(.1305)***	(.1461)***	(.1069)***	(.1041)***
Log Value Added per	.1434	.1362	.0998	.0870
Worker	(.0373)***	(.0377)***	(.0361)***	(.0372)**
Log Input	.1150	.1189	.0722	.0818
	(.0321)***	(.0317)***	(.0304)**	(.0307)***
m1 (Pr>z)	0.000	0.000	0.000	0.000
m2 (Pr>z)	0.251	0.128	0.142	0.141
m3 (Pr>z)	0.127	0.127	0.071	0.065
Sargan	0.000	0.000	0.000	0.000
Hansen	0.000	0.000	0.000	0.000
Number of Instruments	83	53	67	47
Instrument Set	lags 2+	lags 2 and 3	lags 3+	lags 3 and 4
N	3474	3474	3474	3474

Notes: Year dummies and a constant term included in all models. Controls treated as *exogenous*. Robust standard errors in parentheses, clustered at the plant level and adjusted for Windmeijer's correction. See notes to Table 3.