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THE EFFECTS OF FOREIGN MULTINATIONALS ON WORKERS AND FIRMS IN THE UNITED STATES

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ABSTRACT

Governments go to great lengths to attract foreign multinationals because they are thought to raise the wages paid to their employees (direct effects) and to improve outcomes at local domestic firms (indirect effects). We construct the first U.S. employer-employee dataset with foreign ownership information from tax records to measure these direct and indirect effects. We find the average direct effect of a foreign multinational firm on its U.S. workers is a 7 percent increase in wages. This premium is larger for higher skilled workers and for the employees of firms from high GDP per capita countries. We find evidence that it is membership in a multinational production network—instead of foreignness—that generates the foreign firm premium. We leverage the past spatial clustering of foreign-owned firms by country of owner-ship to identify the indirect effects. An expansion in the foreign multinational share of commuting zone employment substantially increases the employment, value added, and—for higher earning workers—wages at local domestic-owned firms. Per job created by a foreign multinational, our estimates suggest annual gains of 13,400 USD to the aggregate wages of local incumbents, two-thirds of which are from indirect effects. Our estimates suggest that—via mega-deals for subsidies from local governments—foreign multinationals are able to extract a sizable fraction of the local surplus they generate.

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1 Introduction

Foreign multinationals account for a sizable fraction of value added, exports, and R&D in the U.S. (BEA, 2017). These firms are affected by regulations on foreign investment, trade policies, and local subsidy competition.[1] A widely held belief is that attracting a foreign multinational to a location will have transformative effects on the outcomes of local workers and firms. The hard evidence on this belief has been limited by data unavailability and the challenge of identifying causal effects. The key questions for policymakers and local stakeholders center on the direct and indirect effects of a job created by a foreign multinational: How much more does a worker earn when she is hired by a foreign multinational? How are domestic firms and their workers in nearby locations affected by foreign firms?

This paper makes four main contributions to understanding the effects of foreign multinationals. First, we use tax records to construct a panel data set for the U.S. that links the population of workers and firms with foreign ownership information of the firms. Second, we develop a model that provides the theoretical underpinnings to study the direct effects that foreign multinationals have on their own workers and the indirect effects that they have on domestic-owned firms and their workers in the local labor market. Third, we leverage the movers between firms to identify the foreign firm premium, i.e., the wage gain for the same worker when moving from a domestic to a foreign firm. Fourth, we document and exploit the spatial clustering of foreign firms to construct an instrument for foreign investment in the local labor market, allowing us to identify the indirect effects of foreign multinationals on the value added, employment, and wages paid at domestic firms.

Our data are created by merging the population of annual U.S. corporate tax filings with the population of annual W-2 tax filings on the wage payments made by employers to workers during 1999-2017.[2] Then, we identify foreign multinationals in these data from a filing requirement for each U.S. corporation that is 25 percent or more foreign-owned. This information also includes the country of foreign ownership. To our knowledge, this is the first paper to combine linked employer-employee panel data with foreign ownership information in the U.S.[3] These panel data provide a unique opportunity to investigate the direct and

[1]The OECD (2019) ranks the U.S. slightly above the OECD average in terms of foreign direct investment restrictiveness. Prominent examples of subsidy deals offered to foreign multinationals include the BMW plant in Spartanburg, South Carolina (1992); the Toyota plant in Blue Springs, Mississippi (2007); and the Foxconn plant in Mount Pleasant, Wisconsin (announced in 2017).

[2]Findings from the matched firm-worker tax records in the U.S. have been reported in studies by Yagan (2019), Kline, Petkova, Williams, and Zidar (2019), Lamadon, Mogstad, and Setzler (2020), and Smith, Yagan, Zidar, and Zwick (2019).

[3]Prior studies on foreign multinationals in the U.S. rely on firm-level data without worker-level information. Several studies combine the Bureau of Economic Analysis (BEA) survey of foreign direct investment in the U.S. and the Census of Manufactures data. Boehm, Flaaen, and Pandalai-Nayar (2019) merge ownership information from LexisNexis Directory of Corporate Affiliations with the Longitudinal Business Database at the Census Bureau. Saha, Firkri, and Marchio (2014) document regional patterns of FDI based on NETS

indirect effects of foreign multinationals in the U.S. labor market.

The primary challenge in studying the direct effects of foreign multinationals on their workers' wages is to disentangle the extent to which higher wages at foreign-owned firms are due to worker skill differentials as opposed to firm premiums. To estimate these premiums, we leverage the U.S. panel data to follow workers who move between foreign and domestic firms. We make four novel contributions to the study of the direct effects of foreign investment. First, this is the first paper to estimate the foreign firm premium in the U.S. that controls for worker skill differentials. We find that the average foreign firm premium is 7 percent. Second, because the U.S. is both the leading headquarter country of multinationals and the top recipient of foreign investment, it provides large samples of both foreign and domestic multinationals. We find that domestic-owned and foreign-owned multinationals have very similar premiums, suggesting that belonging to a multinational network, rather than foreignness, is the main driver of the foreign firm premium. Third, because the U.S. is the top recipient of the world's foreign investment, it provides a rare opportunity to compare the effects of foreign firms by country of origin, with large samples from many diverse countries. We find that the foreign firm premium is increasing in the GDP per capita of the origin country and that firms from higher GDP per capita countries tend to hire more skilled workers. Fourth, it has long been posited that high-skilled workers benefit more from foreign investment, primarily in developing contexts (e.g., Aitken, Harrison, and Lipsey 1996). We provide the first systematic evidence in favor of this hypothesis in the U.S., finding that the wage premium is larger for higher-skilled workers and absent for the lowest decile of worker skill.

Regarding the indirect effects of job creation at foreign firms on local domestic firms and their workers, the key identification challenge is that foreign multinationals may increase employment in a location because of other factors that also cause contemporaneous growth at local domestic firms. To overcome this endogeneity, we document in our data that foreign firms cluster into locations by country of ownership, then exploit this clustering to construct an instrumental variable for local foreign employment.[4] Our identification strategy is analogous to the immigration literature that uses spatial clustering of immigrants to identify the effects of immigrants on native workers' wages (see Card 2001).[5] Equipped with this

data. The data set closest to ours is the one described by the Bureau of Labor Statistics (2019), which has employer-employee links and country of ownership. However, it is for the 2012 cross-section only, and the questions we address in this paper require a panel in order to observe changes over time.

[4] Earlier work by Head, Ries, and Swenson (1995) finds that Japanese affiliates are spatially clustered within the U.S. We are the first to exploit this spatial clustering to identify the indirect effects of foreign multinationals.

[5] While our identification strategy for indirect effects is distinct from the prior literature on spillovers from foreign multinationals, it is more closely related to prior work on agglomeration in urban economics (Bartik, 1991; Moretti, 2010; Combes, Duranton, Gobillon, Puga, and Roux, 2012; Allcott and Keniston, 2018; Helm, 2019).

identification strategy, we find that an increase in employment at foreign-owned firms significantly raises the value added, employment, wage bill, and earnings of continuing workers at domestic-owned firms in the same commuting zone. The effects are larger in the tradable sector than the non-tradable sector and larger among domestic firms with more than 100 employees. Exploring heterogeneity in the wage effects for continuing workers at domestic firms, we find a much larger effect for higher-earning workers and essentially no effect for lower-earning workers. Our estimates imply that, for every 1 job created by a foreign multinational, approximately 0.5 jobs and 139,000 USD in value added are generated at domestic firms in the same local labor market.

With respect to policy implications, our estimates of the direct wage premium by foreign firms highlight sizable benefits of trade and investment policies that encourage foreign firms to invest in the U.S. Furthermore, our estimates imply that local policymakers have incentives to compete for investments by foreign multinationals, for both the direct wage benefits and the sizable local indirect effects on domestic firms and their higher-earning workers. One additional job created by a foreign multinational generates, on average, annual aggregate wage gains for incumbent workers in the commuting zone of approximately 13,400 USD, two-thirds of which are from indirect effects. Outside data suggest that, in the aggregate, foreign multinationals in the U.S. receive 4.6 billion USD in economic development subsidies per year on average.[6] Abstracting from indirect effects, we find that the value of these subsidies is far below the aggregate foreign wage premium of 36 billion USD per year. However, when focusing on the *mega-deals* for large plants, we see that subsidies per job can be quite large. A comparison of our estimates to these subsidy deals reveals that foreign multinationals are able to extract a sizable fraction of the surplus from such investments in the bargaining with local governments for mega-deals. We note that while competing for foreign multinational investments with subsidies may entail local benefits, this does not imply that such subsidies are beneficial from a national welfare perspective; see the discussion by Glaeser and Gottlieb (2009).[7]

The results on direct effects relate to a large existing literature on wage differentials between foreign-owned and domestic-owned firms. Doms and Jensen (1998), Feliciano and Lipsey (1999), and several others find that the average wage at foreign-owned firms is higher

[6]According to data retrieved from the subsidy tracker database of the policy group Good Jobs First, the foreign firm share in total annual economic development subsidies in the U.S. between 2012 and 2017 is about 20 percent. The so-called *mega-deals* (with subsidies larger than 50 million USD) account for about half of all subsidies to foreign firms.

[7]For the analysis of local labor market benefits of various place-based policies, see Gaubert (2018) and Ossa (2017), who model local policymakers using subsidies to compete for firms in spatial equilibrium with agglomeration. Other related studies include business relocation responses to state-level corporate tax changes (Suarez Serrato and Zidar, 2016), agglomeration effects of infrastructure investment (Kline and Moretti, 2013), and indirect effects of employment tax credits (Busso, Gregory, and Kline, 2013).

than that at domestic-owned firms in the U.S. We document in the U.S. tax data that wages are 19 percent higher on average at foreign firms relative to domestic non-multinationals, controlling for observables. Prior studies in other countries have found that the foreign wage premium only explains a small share of the wage differential between foreign-owned and domestic-owned firms (see Heyman, Sjöholm, and Tingvall 2007; Balsvik 2011; and Hijzen, Martins, Schank, and Upward 2013). Our estimate of a 7 percent foreign firm premium implies that two-thirds of the foreign wage differential is the result of worker skill differentials across firms. Thus, the average wage differential shrinks substantially, but is still positive, when accounting for worker skill composition. One possible explanation for the significant wage premium for workers at foreign multinationals is that the U.S. is relatively remote from its major sources of foreign firms (e.g., Europe and Asia), and therefore the selected firms that establish affiliates in the U.S. are especially productive (Helpman, Melitz, and Yeaple, 2004). These firms may also benefit from economies of scale associated with their operations in multiple countries. Another possibility is that firms anchor their wages to headquarter levels, as suggested by Hjort, Li, and Sarsons (2020).

The results on indirect effects relate to a number of studies on productivity spillovers outside the U.S. This literature has found diverse effects. Aitken and Harrison (1999) and Lu, Tao, and Zhu (2017) find negative effects from foreign multinationals on the revenue productivity of domestic firms in the same industry in Venezuela and China, respectively.[8] A number of papers find positive effects on productivity at domestic-owned firms, sometimes associated with buyer-supplier linkages (see Javorcik 2004; Haskel, Pereira, and Slaughter 2007; Alfaro and Chen 2018; Jiang, Keller, Qiu, and Ridley 2018; Kee 2015; Alfaro-Urena, Manelici, and Vasquez 2019b; and Alfaro-Urena, Manelici, and Vasquez 2019a). Poole (2013) finds positive effects on wages at domestic firms from a greater share of coworkers with experience at foreign firms in Brazil, and Driffield and Girma (2003) find that foreign firm entry causes domestic firms to bid up wages. In the U.S. context, Figlio and Blonigen (2000) use variation in foreign investment across counties in South Carolina to find positive effects on county average wages. Analyzing data on publicly traded firms in the U.S., Keller and Yeaple (2009) find positive productivity spillovers from foreign investment on other firms in the same industry.[9] Greenstone, Hornbeck, and Moretti (2010) use a runner-up identification strategy for million-dollar manufacturing plant openings, many of which are owned by multinationals, finding sizable productivity gains for local firms. We contribute to this literature by providing a novel identification strategy for the indirect effects of foreign

[8] Consistent with competition effects, Atkin, Faber, and Gonzalez-Navarro (2018) document a decline in Mexican grocery store prices in response to entry by foreign retailers. See Gorg (2004) for a survey of the empirical literature on FDI spillovers.

[9] Other related work on the indirect effects of foreign multinationals in the U.S. includes Aitken et al. (1996), Branstetter (2001), and Blonigen and Slaughter (2001).

firms and estimating these effects in comprehensive data on workers and firms.

2 Data and Descriptive Evidence

2.1 Data

We now discuss data sources and sample construction; see Appendix A.1 for additional details. We construct a matched worker-firm panel data set from the population of annual U.S. Treasury tax filings from 1999 to 2017. For each worker-firm-year, W-2 tax forms provide information on earnings, the firm's employer identification number (EIN, which is masked to us), and the worker's residential ZIP code.[10] Earnings are defined as all remuneration for labor services deemed taxable by the IRS, including wages and salaries, bonuses, and exercised stock options. We obtain year of birth and sex information from SSA birth records. Following Lamadon et al. (2020), the analysis sample focuses on workers between age 25 and 60 at the highest-paying employment relationship in each worker-year with earnings above the full-time equivalence (FTE) threshold, approximated by the annualized minimum wage.

For each firm-year, Forms 1120 (C-corporations), 1120S (S-corporations), and 1065 (partnerships) provide information on value added and the 6-digit NAICS industry code, where value added equals sales minus cost of goods sold.[11] We refer to the 3-digit NAICS code as the firm's industry and consider the full 6-digit code for robustness.[12] Foreign ownership is indicated by the filing of Form 5472, which is the information return for a U.S. corporation that is 25 percent or more foreign owned and includes the country of foreign ownership. We link worker data to firm data using the EIN. We keep only those firms that have at least one FTE worker. We use the terms "foreign" and "foreign-owned" interchangeably throughout.[13] We consider a firm to be a domestic multinational if it does not file Form 5472 but pays a foreign business tax. Because of difficulties in interpreting value added, we omit the finance, insurance, and real estate (FIRE) industries from all analysis.

[10] In the event that the ZIP code is missing or invalid in year t but not in year s with $|t-s| \leq 2$, and the worker receives a W-2 from the same EIN in t and s, we impute it in t using the value from s.

[11] In manufacturing and mining industries, the cost of goods sold contains production wages (labor compensation to workers directly involved in the production process). We construct a measure of production wages to add back into value added for these sectors (the difference between total wages associated with the firm through worker tax forms and non-production wages reported by the firm).

[12] In the event that the NAICS code is missing or invalid in year t but not in year s with $|t-s| \leq 2$, we impute it in t using the value from s. If this also fails, we impute it from a separate filing, Form 5500.

[13] Similarly, we refer to "domestic" and "domestic-owned" firms interchangeably. We note that even a domestic-owned firm could be in the hands of many small foreign owners, in particular when the company is publicly listed. While we do not have hard data on this, we think these cases are likely to be rare and not necessarily associated with the same effects. In the event that the employer fails to file Form 5472 in year t but files as foreign owned with ownership country c in one of $(t-2, t-1)$ as well as one of $(t+1, t+2)$, we impute foreign ownership in year t as c.

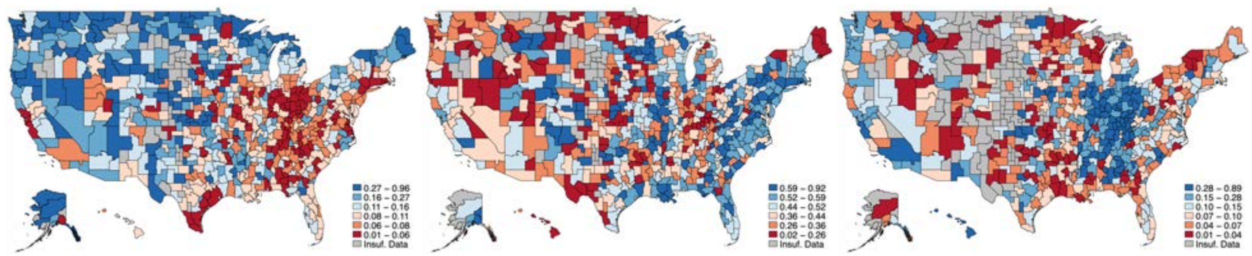
To our knowledge, ours is the first panel data set for the U.S. that links the population of workers and firms with foreign ownership information of the firms. However, working with these data presents two challenges. First, since corporate tax filings provide the foreign ownership information, while the W-2 forms provide the information on employment and wages, we can only classify the foreign status of a worker's firm for those workers whose EIN on the W-2 is also associated with a corporate tax filing. As emphasized by Yagan (2019), many workers cannot be linked to a corporate tax filing, often because the employer is not required to file (especially if the employer is a government or non-profit organization) or because the employer is a subsidiary and only the parent corporation files while the subsidiary uses its distinct EIN to issue W-2 forms. To overcome this challenge, we combine two sources of information on subsidiary linkages. The first source is Schedule K, line 3b, which provides the EIN of the parent corporation in the years in which the subsidiary is a filer, from which we learn the EIN of the parent corporation in future years in which the subsidiary is a non-filer. The second source is the Affiliations Schedule from Form 851, which defines a subsidiary as 80 percent owned by another corporation. However, we only observe a running list of parent-subsidiary relationships taken from the Affiliations Schedules through 2016, so changes over time due to extensive margin changes in subsidiary relationships may be mismeasured when using the second source. For this reason, we only utilize the second source for subsidiary linkages that are not covered by the first source (i.e., subsidiaries that are missing Schedule K filings).

The second challenge is that our analysis requires a firm's activity to be associated with *each* commuting zone in which it is active. This differs from using the address of the firm's headquarter to define its location, as the headquarter may be chosen to obtain favorable state-level tax rates rather than to represent the firm's actual location of activity, and the firm may be active in many locations. Since specific establishments of multi-establishment firms are not observable in U.S. tax data, we follow Yagan (2019) by inferring firms' commuting zone-level operations from workers' residential locations. We aggregate the number of workers and wages within the commuting zone of the worker's address on the W-2 to define the firms' local employment and wage bill. However, we do not observe value added at the firm-commuting zone level directly because it is reported only on firm-level tax forms. To overcome this challenge, we use the share of the wage bill paid in the commuting zone of each firm to allocate value added to commuting zones. For example, if 75 percent of a firm's wage bill is paid in the first commuting zone and 25 percent in the second commuting zone, we allocate 75 percent of value added to the first and 25 percent to the second.

We validate that the data are representative of the share of workers employed by foreign firms using statistics from the BEA and BLS. We find that between 5 and 6 percent of American workers are employed at foreign firms and the average worker at a foreign firm earns

Figure 1: Geographic Clustering of Foreign Firms by Country of Origin

(a) Owned in Canada (b) Owned in Western Europe (c) Owned in East Asia



Notes: The figures display spatial variation in the concentration of foreign employment that is at firms owned in particular groups of owner countries based on total FTE worker-year observations from 1999-2017.

25 percent more than the average worker at a domestic firm, which match the statistics from the Bureau of Labor Statistics (2019). Appendix Figure A.1 visualizes the share of American workers employed at foreign firms between 1977 and 2017. It compares three series available from the BEA to the series we construct from tax data. Each series follows different sample selection rules, yet during the years of overlap, the series are generally consistent. This figure also illustrates the striking rise in the importance of foreign-owned firms in the U.S. labor market. Only 2 percent of workers were employed by foreign-owned firms in the late 1970s, whereas around 6 percent are employed by foreign-owned firms today.

2.2 Descriptive Statistics

Differences between foreign and domestic firms. Appendix Table A1 provides summary statistics on foreign and domestic firms for the year 2015. Four clear differences can be noted. First, the average foreign firm operates in about 7 locations, whereas the average domestic firm operates in about 2 locations. Second, the average foreign firm is much larger than the average domestic firm, with about 28 workers per domestic firm and 172 workers per foreign firm. Third, value added per worker in the analysis sample is 220,100 USD at foreign firms and 153,100 USD at domestic firms, indicating that value added per worker is more than 40 percent higher at foreign firms. Fourth, the average worker in the analysis sample earns 75,700 USD at foreign firms and 60,700 USD at domestic firms, indicating 25 percent higher wages at foreign firms.[14]

[14]Relatedly, Appendix Figure A.2 provides value added and wage differentials (relative to the average domestic non-multinational firm) by country of origin for the 34 countries with the most unique firms operating in the U.S. during 2010 to 2015. Specifically, we select the 40 countries with the most firms in 2010-2015 and drop five tax haven countries (e.g., Cayman Islands) as well as the "other country" category. We see a clear pattern that the value added and wage differentials between foreign and domestic firms are greater for countries of origin with higher GDP per capita.

Spatial distribution of foreign employment. In Appendix Figure A.3a, we plot the share of workers employed at foreign firms in 2001 for each commuting zone. We find particularly high levels of employment at foreign firms along the East Coast and in Rust Belt cities in Indiana, Michigan, and Ohio, but especially low levels in the South. In Appendix Figure A.3b, we illustrate the changes in the share of employment at foreign firms by commuting zone from 2001 to 2015. Substantial changes have taken place across the U.S., with Gulf Coast states such as Alabama, Louisiana, and Mississippi experiencing especially rapid growth, while parts of the East Coast and the Rust Belt have experienced sharp declines in the share of foreign employment.

Clustering by nationality. In Figures 1a, 1b, and 1c, we display the share of employment at Canadian, Western European, and East Asian firms as a share of total employment at foreign-owned firms by commuting zone, based on total FTE worker-year observations from our sample. A clear visual pattern emerges: Canadian firms are more likely to be near the Canadian border, European firms are primarily engaged in the eastern part of the U.S., and Asian firms account for a large share of foreign-owned firms near the West Coast as well as in the Midwest.

There are a number of plausible reasons why firms cluster by nationality. First, the cost of shipping intermediate goods from the home country or the costs of communication may lead to clustering on distance (Keller and Yeaple, 2013) or clustering on the availability of airline routes to the headquarter (Giroud, 2013; Campante and Yanagizawa-Drott, 2018). Second, foreign firms may be more likely to hire employees (in particular, managers) from their country of origin that already had business experience at the firm's headquarter, who may prefer to live near other immigrants from their country.[15] Third, foreign firms of a particular country of origin may share information, for example, by using similar plant site selection firms that already have business and political contacts in certain regions. Fourth, firms may cluster by industry, and some countries specialize in particular industries (Head et al., 1995). This clustering by country of ownership will be important when discussing our identification strategy for indirect effects in Section 5.

3 A Model of Foreign Multinationals

In this section, we develop a model in which foreign multinationals pay wages that are different from those of domestic firms to a worker of a given skill type (direct effects) and affect outcomes at local domestic firms (indirect effects). Rather than from foreignness per se,

[15]Relatedly, Burchardi, Chaney, and Hassan (2019) document for the U.S. that foreign investment follows past ancestors' regional choices.

direct wage effects arise because more productive firms need to pay higher wages to recruit their marginal employee. Furthermore, firms belonging to a multinational network may have access to more skill-augmenting technology, leading them to disproportionately employ higher-skilled labor and pay a greater premium to higher-skilled labor relative to lower-skilled labor. Indirect effects can arise from technology spillovers – which are beneficial to domestic firms – and competition effects – which are harmful to domestic firms. For brevity, the main text develops the case with two skill types (skilled and unskilled) and two firm nationalities (foreign and domestic). Appendix A.3 provides derivations. Appendix A.4 provides a more general case with an arbitrary number of skill types and firm types that differ by country of origin.

3.1 Model

Environment. We assume there is a large set of locations in the U.S. All regions are trading frictionless within the U.S., and workers are immobile across locations. We focus on the outcomes in one particular location and, to simplify notation, omit the location subscript. Let $N \in \{D, F\}$ denote the firm country of origin, where D is domestic and F is foreign. Denote by M_N the number of firms of nationality N. Let $h \in \{s, u\}$ denote the skill type of a worker, where s denotes skilled and u denotes unskilled. Denote by L_{Nh} the number of employees at firms from nationality N with skill level h, and $L_N = \sum_h L_{Nh}$ is the total number of employees for nationality N. The share of workers that are skilled in a nationality N firm is $C_N \equiv \frac{L_{Ns}}{L_N}$. Each region is equipped with $\bar{L}_h$ potential employees of skill type h, and the employment rate is $E_h \equiv \frac{\sum_N L_{Nh}}{\bar{L}_h}$. In each location, the composition of skilled workers by nationality, C_N, as well as the local employment rate, E_h, are equilibrium objects.

Technology. Each firm produces a homogeneous good q that is freely traded, where the price is normalized to 1. A firm of nationality $N \in \{D, F\}$ produces using technology,

$$q_N(\ell_u, \ell_s) = \phi_N \left(\ell_u + \zeta_{Ns}\ell_s\right) \tag{1}$$

where ϕ_N is total factor productivity (TFP) and ζ_{Ns} is skilled labor augmenting productivity. We assume, and later provide evidence, that foreign firms are more productive than domestic firms in their usage of both unskilled labor (i.e., $\phi_F \geq \phi_D$) and skilled labor (i.e., $\phi_F\zeta_{Fs} \geq \phi_D\zeta_{Ds}$). Helpman et al. (2004) provide a micro-foundation in which foreign firms are more productive because they must overcome a larger fixed cost of entry. While we take the TFP of foreign firms $\phi_F > 1$ as determined prior to market entry, we allow for spillovers of TFP from foreign to domestic firms as

$$\phi_D = 1 + \tau \frac{L_F}{L_D + L_F}(\phi_F - 1), \tag{2}$$

where $0 \leq \tau \leq 1$ is the spillover rate. When $\frac{L_F}{L_D+L_F}$ is greater, domestic firms are more exposed to foreign multinationals, and τ determines the sensitivity to this exposure.

Labor supply. Let w_{jh} denote the wage offered by firm j to a worker of skill type h. The utility of worker i when employed at a given firm j with wage offer w_{jh} is,

$$V_{ij} = \log w_{jh(i)} + \epsilon_{ij}, \tag{3}$$

where the wage of the outside option (non-employment) is w_0. Unobserved preferences ϵ_{ij} can be determined by a wide range of characteristics, such as distance of the firm from the worker's home. Following recent work by Card, Cardoso, Heining, and Kline (2018), Lamadon et al. (2020), and Berger, Herkenhoff, and Mongey (2019), we parameterize ϵ_{ij} as i.i.d. type 1 extreme value with dispersion $1/\eta$. When ϵ_{ij} is more dispersed (i.e., $1/\eta$ is greater), our preference specification allows workers to view firms as worse substitutes. Letting ℓ_{jh} denote the number of workers of skill type h in firm j, the implied labor supply to firm j is

$$\ell_{jh} = w_{jh}^{\eta} \frac{\bar{L}_h}{W_h}, \tag{4}$$

where $W_h = \sum_{k=0}^{M_D+M_F} w_{kh}^{\eta}$ is the aggregate wage index. Equation (4) shows that η can be interpreted as the firm-specific labor supply elasticity.

Labor demand. Since ϵ_{ij} is unobserved to the firm, firms cannot price discriminate on idiosyncratic preferences and thus post a common wage for all workers of skill type h. We assume that there are many firms of its type in its region, so each firm acts monopsonistically competitive, meaning it does not take the effect of its own choice of w_{jh} or ℓ_{jh} on W_h into account. Given the production function in equation (1) and labor supply in equation (4), and normalizing $\zeta_{Nh} = 1$ for $h = u$, a firm with nationality N offers wage

$$w_{Nh} = \frac{\eta}{\eta + 1} \phi_N \zeta_{Nh} \quad N \in \{D, F\}, \quad h \in \{s, u\}. \tag{5}$$

Since $\phi_N \zeta_{Nh}$ is the marginal product of labor for skill type h at a firm of nationality N, $\frac{\eta}{\eta+1}$ is the markdown on the marginal product of labor.

3.2 Direct effects

From equation (5), the mean difference in log wages between foreign and domestic firms is

$$\underbrace{\mathbb{E}\left[\log w_{F\cdot}\right] - \mathbb{E}\left[\log w_{D\cdot}\right]}_{\text{Total foreign wage differential}} = \underbrace{\log \phi_F - \log \phi_D}_{\text{Unskilled foreign firm premium}} + \underbrace{C_F \log \zeta_{Fs} - C_D \log \zeta_{Ds}}_{\text{Composition-weighted skilled foreign firm premium}} . \tag{6}$$

In the absence of skill-augmenting technology ($\zeta_{Fs} = \zeta_{Ds} = 1$), skill composition is the same in foreign and domestic firms ($C_F = C_D$), so the total foreign wage differential simplifies to the productivity difference ($\log \phi_F - \log \phi_D$). For the more interesting case in which technology is skill-augmenting, we summarize equation (6) with the following proposition:

Proposition 1 (Direct effects) *If the TFP of foreign firms is greater than domestic firms (i.e., $\phi_F > \phi_D \geq 1$) and the production technology at foreign firms is more skill-augmenting relative to domestic firms (i.e., $\zeta_{Fs} > \zeta_{Ds} \geq 1$), then*

(a) The unskilled foreign firm premium is positive;

(b) The skilled foreign firm premium is greater than the unskilled foreign firm premium;

(c) The skill composition is greater at foreign firms (i.e., $C_F > C_D$).

3.3 Indirect effects

We next investigate the indirect effects (i.e., the effects of entry and expansions by foreign firms on domestic firms). Because of the complex nature of the model, our focus is on providing the predicted effects of foreign firm entry based on first-order approximations. Let $\Delta y \equiv y' - y$ denote a change to y. The effects of interest center on $\hat{X} \equiv \frac{\Delta L_F}{L_D + L_F}$, which is a small perturbation in employment at foreign firms relative to initial employment at all firms, and we take the initial equilibrium to feature a small share of employment at foreign firms when deriving the first-order approximation of equilibrium outcomes.

Wage. A first-order approximation of the wage at domestic firms yields the prediction

$$\underbrace{\Delta \log(w_{Dh})}_{\text{Domestic firm wage change}} \approx \underbrace{\tau(\phi_F - 1)\hat{X}}_{\text{Technology spillover effect}} . \tag{7}$$

This equation states that the wage increase at domestic firms is proportional to the TFP increase at domestic firms.[16] The magnitude of the TFP increase depends on the spillover

[16] We show in Appendix A.3 that this prediction does not rely on the first-order approximation and holds more generally (i.e., $\frac{dw_{Dh}}{dM_F} > 0$ if $\tau > 0$ and $\frac{dw_{Dh}}{dM_F} = 0$ if $\tau = 0$).

rate τ, the relative productivity of foreign firms $\phi_F - 1$, and the relative size of entering foreign firms $\hat{X}$.

Employment. Let $\bar{E}_N \equiv C_N E_s + (1 - C_N) E_u$ denote the nationality skill composition-weighted average labor market tightness. A first-order approximation for employment at a domestic firm is

$$\underbrace{\Delta \log(\ell_{Du} + \ell_{Ds})}_{\text{Domestic firm employment change}} \approx \underbrace{\tau\eta(\phi_F - 1)\left(1 - \bar{E}_D\right)\hat{X}}_{\text{Technology spillover effect}} - \underbrace{\bar{E}_F \hat{X}}_{\text{Competition effect}}. \tag{8}$$

The equation shows that the employment response at domestic firms can range from negative to positive. Because of labor market competition effects, the model *without* productivity spillovers (i.e., $\tau = 0$) implies a *decline* in the output at domestic firms as the activity by foreign firms in a location increases. With large enough productivity spillovers, employment at domestic firms increases when the employment share at foreign firms grows. If the labor market is less tight (lower $\bar{E}$) or labor supply is more elastic (higher η), the technology spillover effect becomes stronger. Furthermore, competition effects are weaker when the labor market is less tight.

Value added and wage bill. Denote by $R_N \equiv \frac{\zeta_{Ns}\ell_{Ns}}{\ell_{Nu}+\zeta_{Ns}\ell_{Ns}}$ the share of output at a firm with nationality N that is produced by skilled workers. The object R_N differs from C_N in that it depends on the skill-augmenting productivity ζ_{Ns}. Using a first-order approximation,

$$\begin{aligned}\underbrace{\Delta \log q_D}_{\text{Domestic firm value-added change}} &\approx \underbrace{\tau(\phi_F - 1)\left(1 + \eta\left[1 - R_D E_s - (1 - R_D)E_u\right]\right)\hat{X}}_{\text{Technology spillover effect}} \\ &\quad - \underbrace{\left(\frac{C_F}{C_D} R_D E_s + \frac{1 - C_F}{1 - C_D}(1 - R_D)E_u\right)\hat{X}}_{\text{Competition effect}}.\end{aligned} \tag{9}$$

Since the value added and wage bill are proportional, equation (9) is also the first-order approximation to the log change in the wage bill. Similar to the employment response, the change in value added at domestic firms can range from negative to positive, depending on the same set of factors as the employment response but also depending on R_D, C_D, and C_F.

Value added per worker. Whether the log value added response (equation 9) exceeds the log employment response (equation 8) at domestic firms, and hence value added per worker increases, turns on various factors. In the simple case in which skilled and unskilled labor are symmetric (i.e., $C_F = C_D$, $E_s = E_u$, and $R_D = 0.5$), value added per worker must increase in the presence of technology spillovers in response to foreign firm entry. However, if foreign firms are more skill intensive ($\frac{C_F}{C_D} > 1$), an expansion in employment at foreign firms leads domestic firms to substitute toward unskilled labor. All else equal, the substitution toward unskilled labor lowers value added per worker at domestic firms. Therefore, value added per worker at domestic firms could decrease even in the presence of positive technology spillovers. A similar argument holds for wage bill per worker—unskilled workers receive lower wages, so substitution toward unskilled labor lowers wage bill per worker, all else equal.[17]

We summarize the above indirect effect predictions in a proposition:

Proposition 2 (Indirect effects) *If the TFP of foreign firms is greater than domestic firms* ($\phi_F > \phi_D \geq 1$) *and foreign firms have positive spillovers onto domestic firms (i.e.,* $\tau > 0$*), then — up to a first-order approximation around an initial equilibrium featuring a small share of employment at foreign firms — an increase in the share of employment at foreign firms causes*

(a) A positive effect on wages at domestic firms;

(b) A positive effect on employment, wage bill, and value added at domestic firms if $\tau(\phi_F - 1)$ *is sufficiently large or* E_s *and* E_u *are sufficiently small;*

(c) Ambiguous effects on value added per worker and wage bill per worker at domestic firms.

3.4 Model extensions and limitations

Before proceeding to the empirics, we note several limitations of the model. Clearly, the model is highly stylized with only two types of workers and two types of firms. In Appendix A.4, we provide a more general case with an arbitrary number of skill types and firm types that differ by country of origin (where firms from different countries of origin can have access to different technologies). Regarding direct effects, our model predicts that firms from countries of origin with more skill-augmenting technology will disproportionately employ higher-skilled labor and pay a greater premium to higher-skilled labor relative to lower-skilled labor. We confirm this prediction in the next section when estimating a wage model with many skill types and many firm types.

By assuming that output is freely tradable, the model abstracts away from the product market competition effects associated with foreign firm entry in a commuting zone. See Bloom, Schankerman, and Van Reenen (2013) for a method of separating product market competition effects from technology spillover effects. Furthermore, the simple model abstracts away from input-output linkages between firms. Access to cheaper local inputs

[17] For this reason, it is preferred to measure the indirect effects on wages (equation 7) using continuing workers rather than wage bill per worker in the empirical application below.

or an increase in local demand would affect domestic firms' outcomes in a similar way as technological spillovers.

4 Direct Effects of Foreign Multinationals

We next empirically examine the direct effects of foreign multinationals on workers in the U.S. Our primary goal is to disentangle the extent to which higher wages at foreign-owned firms are due to worker skill differentials as opposed to firm premiums. We leverage the U.S. data to make four novel contributions about the direct effects of foreign investment. First, this is the first paper to estimate the foreign firm premium in the U.S. that controls for worker skill differentials. We find that the average foreign firm premium is 7 percent. Second, because the U.S. is both the leading headquarter country of multinationals and the top recipient of foreign investment, it provides large samples of both foreign and domestic multinationals. We provide the novel finding that domestic-owned and foreign-owned multinationals have very similar premiums. Third, because the U.S. is the top recipient of the world's foreign investment, it provides a rare opportunity to compare the effects of foreign firms by country of origin, with large samples from many diverse countries. We reach the novel finding that the foreign firm premium is increasing in the GDP per capita of the origin country. Fourth, it has long been posited that high-skilled workers benefit more from foreign investment, primarily in developing contexts (e.g., Aitken et al. 1996). We provide the first systematic evidence in favor of this hypothesis in the U.S.

4.1 Estimation Strategy for the Foreign Firm Premium

We now consider estimating the equilibrium wage equation (5) from Section 3, but with the extension derived in Appendix A.4 to allow for an arbitrary number of firm and worker types. For simplicity, we initially restrict the skill-augmenting technology parameter to be constant across firms. Under this restriction, the equilibrium wage setting with many skill and firm types is[18]

$$\log w_{i,t} = \psi_{j(i,t)} + x_i + \chi'_{i,t}\beta + \epsilon_{i,t}, \tag{10}$$

where $j(i,t)$ denotes firm j that employs worker i in year t, ψ denotes the firm premium, x denotes worker skill, and χ denotes a vector of observable determinants of earnings.[19] Our main specification estimates equation (10) for years 2010-2015 on the largest connected set

[18]The derivation of equation (10) is provided in Appendix A.4 without ϵ. We include the idiosyncratic unobservable ϵ in the empirical implementation to allow for measurement error. We provide estimates when allowing for heterogeneous skill-augmenting productivity parameters in Subsection 4.3.

[19]Song, Price, Guvenen, Bloom, and von Wachter (2018) and Lamadon et al. (2020) also estimate (10) on the U.S. tax data, but do not examine foreign ownership.

of firms, with robustness checks presented below.[20] In χ, we control for location-year fixed effects, industry-year fixed effects, and a third-order polynomial in the age of the worker.

Our aim is to estimate equation (10) to characterize differences in ψ and x across countries of ownership. Equation (10) is identical to the two-way fixed effects regression proposed by Abowd, Kramarz, and Margolis (1999). The key identifying assumption for this regression is that workers do not select to move into firms based on the idiosyncratic error ϵ. However, selection based on the worker effects x, firm effects ψ, or observable controls χ does not violate identification. Card et al. (2018) propose an event study representation to visualize potential selection on ϵ. If the log wage residuals (controlling for χ) are on different trends for those who move into different firm types, this suggests workers select on ϵ, as x and ψ are time invariant. Since our goal is to identify the premium for foreign versus (non-multinational) domestic firms, we consider an analogous event study for workers who move between foreign and domestic firms in Appendix A.5. As demonstrated in Appendix Figure A.4, there is little evidence of pre-trends prior to the moves, which is consistent with a measurement error interpretation of ϵ. Furthermore, when restricting to the sample of workers who lose their jobs in a mass layoff (and therefore are even less likely to select to move based on individual-specific idiosyncratic errors), pre-trends are virtually the same as in the full sample.[21]

An important difficulty in estimating equation (10) remains. As shown by Andrews, Gill, Schank, and Upward (2008), *limited mobility* makes it challenging to precisely estimate firm premiums and worker effects. The earnings changes for workers who move across firms provide the identifying content on firm premiums, and the bias in those firm premium estimates declines as the number of movers per firm grows. However, the modal firm in the U.S. has *a single mover*, providing the opportunity for massive limited mobility bias in our context. To address this, we follow the approach of Bonhomme, Lamadon, and Manresa (2019) and estimate a set of grouped fixed effect models. Instead of obtaining a fixed effect for each firm, we allocate all firms in our data to $k = 10$ clusters ($k = 20, 30, 40, 50$ in robustness checks) with similar wage structures using k-means cluster analysis.[22] These clusters preserve the wage structure while reducing the number of fixed effects that must be estimated. Indeed, we find that 86 (92) percent of all between firm earnings variance is captured by only these 10 (50) clusters. Since there is much more mobility between these clusters than between the millions of unique firms, any bias should be mitigated. Lastly, by providing a parsimo-

[20]Equation (10) is typically estimated on short time intervals, as fixed effects are a worse approximation to the wage structure over a longer period of time (see the discussion by Lamadon et al. 2020 and Lachowska, Mas, Saggio, and Woodbury 2020).

[21]We follow Yagan (2019) in using a 30 percent separation rate to define a mass layoff event.

[22]Lamadon et al. (2020) are the first to provide bias-corrected estimates of firm premiums and sorting for the U.S. Using the grouped fixed effects approach, they find that the variance of firm premiums is inflated by a factor of about three when ignoring limited mobility bias, while the correlation between worker skill and firm premiums is deflated by a factor of about four.

nious representation of firm heterogeneity, the k-means clustering procedure will also make it feasible below to estimate the more general model in which skill-augmenting productivity parameters are heterogeneous across firm types and, therefore, workers of different skill levels receive different premiums.

4.2 Main Results on Foreign Firm Premiums

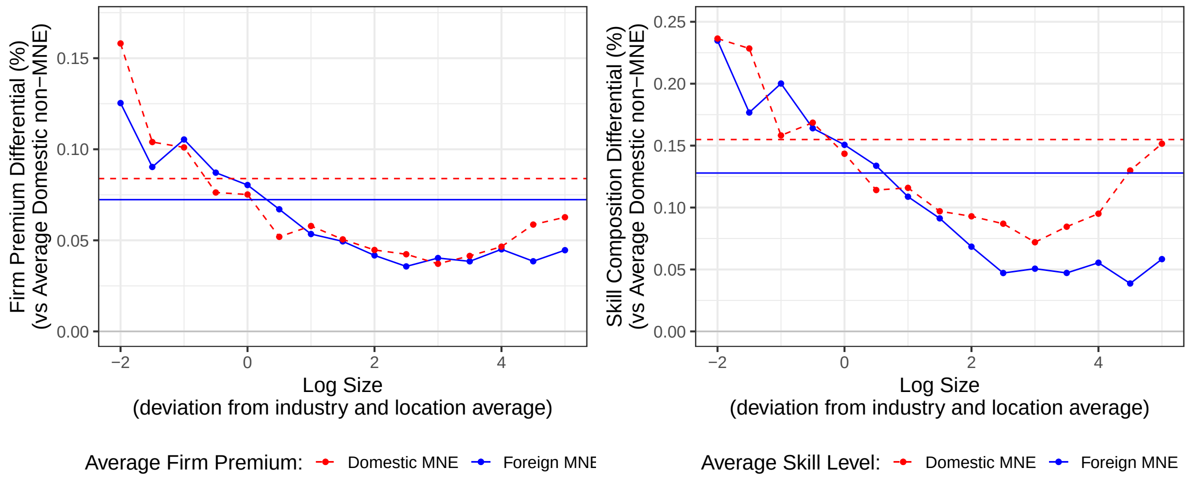
We now provide the main estimates from equation (10). Throughout the analysis, we take domestic non-multinationals as the reference group of firms. We treat domestic multinationals as a distinct group of firms so that we can investigate the similarity between domestic and foreign multinationals. Controlling for the observables listed above, the average worker at a foreign multinational earns 19.5 percent more than the average worker at a domestic non-multinational, while workers at domestic multinationals earn 23.0 percent more on average. Using the estimates based on equation (10), we find that the average firm premium is 7.2 percent at foreign multinationals and about 8.4 percent at domestic multinationals. From the decomposition in equation (6) (and the analogous expression with many skill and firm types in Appendix A.4), this indicates that, at both foreign and domestic multinationals, about two-thirds of the residual wage differential is due to a greater composition of high-skill workers at foreign multinationals relative to domestic non-multinationals. Recall that we control for industry-year and commuting-zone-year fixed effects in all direct effects estimation, so reported differentials in log earnings, firm premiums, and worker composition do not reflect location or industry selection.

In Figure 2, we show that the average firm premiums and worker compositions of foreign and domestic multinationals track one another closely across the firm size distribution. This evidence suggests that belonging to a multinational network, rather than foreignness, is the main driver of the foreign firm premium. Multinational firms are more productive through selection—it is the most productive firms that can overcome the entry costs to establish foreign affiliates (Helpman et al., 2004). Furthermore, belonging to a multinational network confers productivity advantages through access to additional sources of inputs and technology. An implication is that domestic and foreign multinationals are expected to be more productive and thus have substantial firm premiums relative to the reference group. Relatedly, Bloom and Van Reenen (2007, 2010) find that management practices are similar and of high quality for multinational relative to domestic firms across countries of ownership.

Both domestic and foreign multinationals hire more skilled workers compared to non-multinationals. The difference in the skill composition holds across the entire firm size distribution but is particularly pronounced when comparing a smaller multinational firm to a non-multinational firm in the same size bin (Figure 2b). Similarly, the multinational wage premium appears to be highest when comparing between smaller firms (Figure 2a). The

Figure 2: Comparison of Foreign and Domestic Multinationals

(a) Firm Premium Differential (%) (b) Skill Composition Differential (%)



Notes: This figure presents estimates of the model in equation (10) from the grouped fixed effect estimator during 2010-2015. The horizontal axis is an equally-spaced grid of width 0.5 in the residual log firm size distribution, where each unit is associated with the nearest grid point. The vertical axis is the difference in the average firm premium (subfigure a) or average worker skill level (subfigure b) for foreign (blue) or domestic (red) multinationals, relative to the average domestic non-multinational in the same size bin. The horizontal lines indicate the overall averages (not conditional on a size bin).

similar shape of the multinational wage premium and skill composition differences across the size distribution may be related. In Section 4.3, we explore the extent to which the wage premium differs by worker skill type.

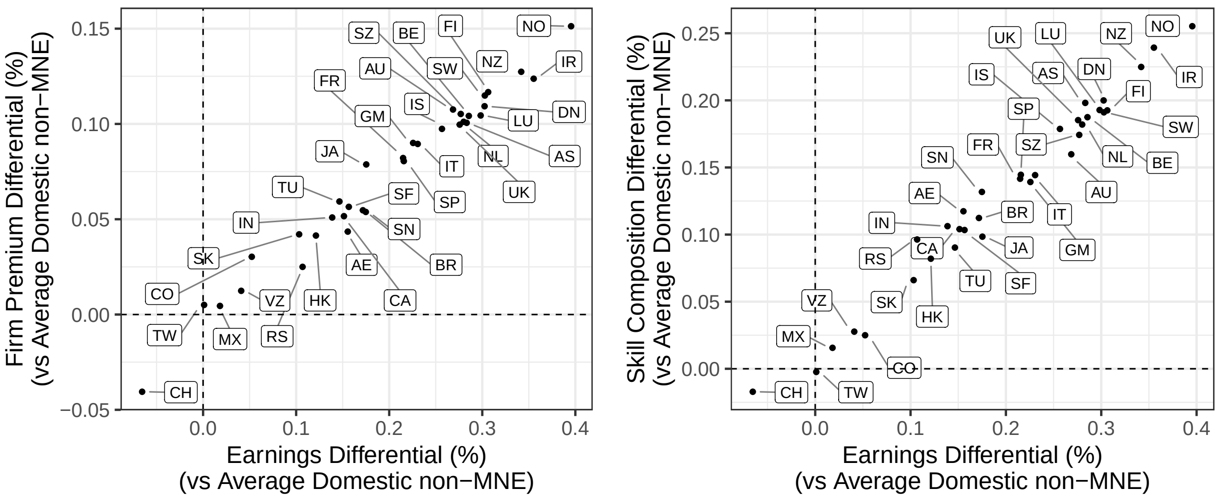
An important feature of the U.S. data is that there are sufficiently many unique firms from a large number of foreign countries to estimate country-specific foreign firm premiums. Figure 3(a) plots the mean firm premium estimate for the 34 countries of ownership with the most firms against mean log earnings, where mean log earnings is normalized to be zero at domestic non-multinational firms. We find substantial heterogeneity in the firm premium by country of origin. The Northern European countries of Norway, Finland, Sweden, and Denmark, as well as Ireland and New Zealand, have larger than average firm premiums. At the other extreme, small positive firm premiums are estimated for Colombia, Mexico, Russia, Taiwan, and Venezuela, while a *negative* 4 percent premium is estimated for China. The share of the wage differential explained by firm premiums is approximately the same across all countries at around 37 percent. This means that countries that offer higher premiums also attract more talented workers, as shown in Figure 3(b).

There are many possible reasons for this heterogeneity across countries of ownership. As the cost of entry increases, we expect the average premium of entering firms to increase.[23]

[23]Distance is a suggested mechanism by Helpman et al. (2004). Egger, Jahn, and Kreickemeier (2018) find

Figure 3: Foreign Firm Premiums and Worker Skill Composition

(a) Firm Premium Differential (%) (b) Skill Composition Differential (%)



Notes: This figure presents estimates of equation (10) from the grouped-fixed effect estimator during 2010-2015. The vertical axis is the difference in the average firm premium (subfigure a) or average worker skill level (subfigure b) for foreign multinationals with the countries of ownership indicated by the labels, relative to the average domestic non-multinational.

Another possibility is that firms anchor their wages to headquarter levels, as suggested by Hjort et al. (2020). Finally, it could be that countries with greater GDP per capita have access to more skill-augmenting technology (Caselli and Coleman, 2006), which could explain higher firm premiums (we explore this case analytically in the model of Appendix A.4). To investigate this issue, Appendix Figure A.6(a) plots the mean firm premium estimate for these countries of ownership against log GDP per capita, observing a clear pattern that countries of ownership with higher GDP per capita provide greater average premiums to their workers. Regressing the average firm premium on log GDP per capita and log distance from the U.S. yields a highly statistically significant coefficient of 0.031 for log GDP per capita and a statistically insignificant coefficient of 0.011 for log distance. This suggests that GDP per capita is more important than distance in explaining country heterogeneity in the firm premium. We find a similar pattern for average skill composition by GDP per capita in Appendix Figure A.6(b). These findings are consistent with countries with higher GDP per capita having access to more skill-augmenting technology, leading to a higher composition of skilled workers and greater premiums as GDP per capita rises.

a pattern of foreign firm wage differentials that increase in distance to the headquarter country in Germany.

4.3 Extension to Allow for Skill-augmenting Productivity

The model of Section 3 allows for skill-augmenting productivity to differ between foreign and domestic firms. In Appendix A.4, we generalize this model to allow for an arbitrary number of firm and worker types, which yields a more general regression,

$$\log w_{i,t} = \psi_{j(i,t)} + \theta_j x_i + \chi_{i,t}'\beta + \epsilon_{i,t}, \tag{11}$$

where, if θ_j is greater at foreign relative to domestic firms on average, then foreign multinationals have more skill-augmenting technology and in turn pay a greater relative premium to high-skilled workers.[24] Bonhomme et al. (2019) provide a method for estimating equation (11); for brevity, we review main results here while providing a detailed explanation of the estimator and findings in Appendix A.6.

We find that the foreign firm premium is monotonically increasing in the skill of workers compared to the premium offered by domestic non-multinationals to workers of the same skill level. Foreign multinationals pay a 19 percent greater premium to workers in the top skill decile, but a 1 percent negative premium to workers in the bottom skill decile. Furthermore, we find that domestic-owned multinationals pay a 21 percent greater premium to workers in the top skill decile than domestic non-multinationals, but no premium to workers in the bottom skill decile. These results are consistent with multinationals having more skill-augmenting technology than non-multinationals. Skill-augmenting technology would lead multinational firms (both foreign-owned and domestic-owned) to bid up the price of local labor for skilled workers such as managers, as found by Bloom, Brynjolfsson, Foster, Jarmin, Patnaik, Saporta-Eksten, and Van Reenen (2019), but not bid up the price of routine labor.

4.4 Robustness of the Foreign Firm Premium Estimates

Our main estimate of the average foreign firm premium is robust to various alternative specifications. First, the grouped fixed effects estimator of equation (10) requires specifying the number of clusters to use in the k-means algorithm. Appendix Figure A.7 demonstrates that the results are nearly identical when allowing for 10, 20, 30, 40, or 50 clusters, with an average foreign firm premium of 7.2 percent relative to domestic non-multinational firms in each case. Second, we find that the results are robust to controlling for third-order polynomials in log firm size (with polynomials in both the firm's local employment and national employment across all of the firm's locations), with a mean foreign firm premium estimate of 6.2 percent. Third, Appendix Figure A.8 demonstrates that the results are nearly

[24] Note that equation (10) is the special case of equation (11) in which $\theta_j = \bar{\theta}, \forall j$, that is, the skill-augmenting productivity is homogenous. Equation (11) was estimated in the U.S. by Lamadon et al. (2020), who also find evidence that θ_j varies across firms, but they do not examine foreign ownership.

the same when performing the estimation for the 2001-2006 sample rather than the 2010-2015 sample considered above, with an average foreign firm premium of 6.7 percent relative to domestic non-multinational firms in 2001-2006. Fourth, when allowing for firm-worker interactions as discussed above, the average foreign firm premium is 7.8 percent on average relative to domestic firms. Fifth, in Appendix A.5, we use a difference-in-differences design for workers that move across firms as a distinct but complementary approach to equation (10).[25] As reported in Appendix Table A2, we find that moving between domestic and foreign firms is associated with a 5-8 percent wage change (relative to wage growth for workers who move between domestic firms), which is similar to the main estimate. The estimates are in the 5-6 percent range when considering only moves that occurred in a mass layoff event at the worker's initial employer.

4.5 Mechanisms behind the foreign firm premium.

We briefly discuss five alternative explanations for the foreign firm premium. We do not find any as convincing as the productivity selection mechanism of Helpman et al. (2004).

Hours. One possibility is that the same worker earns more at a foreign firm because of working longer hours. While the tax data do not include information about hourly wages, according to survey data by the Bureau of Labor Statistics (2019), foreign firms pay 20 percent more than domestic firms even for workers in production occupations for which the reported wages should be primarily at the hourly wage instead of the annual salary level.[26] We therefore think it is unlikely that hours worked are the main driver of foreign premiums.

Layoff risk. Foreign firms may be perceived as being more risky employers, as existing research has found (domestic) multinationals to be at greater risk of shutting down plants than non-multinational firms of similar size (Bernard and Jensen 2007). However, plant shutdowns account for only a small fraction of overall job separations. We find that the probability of staying at the same employer next year is actually higher for workers at foreign firms than for workers at domestic firms. We also find a lower likelihood of separations due to mass layoffs at foreign firms (see the sample sizes in Appendix Table A2). Therefore, the risk of job separation – both overall and due to layoffs – appears to be lower at foreign firms.

[25] An advantage of this approach is that it is straightforward to visualize the pre-trends, as discussed in Subsection 4.1. A disadvantage is that it does not yield the joint distribution of (ψ, x) needed for the various dimensions of heterogeneity we explore.

[26] According to the Current Population Survey, 80 percent of workers in production occupations receive hourly wages as opposed to a fixed annual salary. The instructions in the Occupational Employment Report ask firms to report hourly wages for part-time workers as well as for salaried workers, who do not work a standard 2,080 hours per year (40 hours per week).

Amenities and fringe benefits. It could be that foreign firms have lower amenities than domestic firms, and thus must be paid greater wages to achieve a similar level of compensation. We have not been able to find systematic data on this claim. Anecdotes, however, suggest that foreign firms tend to be attractive employers overall. Examining the 20 employers ranked as having the "Top 20 Employee Benefits and Perks for 2017" in the U.S. by Glassdoor, we see that 5 (25 percent) are foreign owned.[27] In survey data from Costa Rica, Alfaro-Urena et al. (2019a) find that amenities and fringe benefits are better at foreign-owned firms.

Stigma. A stigma may be associated with working at a foreign-owned firm, for which higher wages compensate. While such a stigma may exist, our evidence presented in Figure 3 shows that the wage premium is rising with GDP per capita of the owner country, whereas we might expect stigma to be negatively associated with GDP per capita of the owner country.

Information or monitoring costs. Foreign owners may have worse information about the skill of the workers they hire and overpay them. Alternatively, monitoring workers may be more difficult for foreign owners (Head and Ries, 2008). In lieu of monitoring, firms may pay a premium to discourage workers from shirking, and the premium may be greater for workers with greater ability or those in positions of responsibility (Oi, 1983; Katz, 1986). We note that it would not affect our conclusion of a positive effect of foreign firms on their workers if the premium were due to information or monitoring costs.

5 Indirect Effects of Foreign Multinationals

As discussed in Section 3, in addition to directly affecting the wages of their own workers, foreign multinationals may also affect domestic firms and their workers indirectly. The theory suggests that these effects can be positive or negative.

5.1 Empirical Strategy to Estimate Indirect Effects

In this section, we seek to measure the indirect effects of employment growth at foreign-owned firms on outcomes at domestic-owned firms. Using a functional form suggested by the first-order approximations derived in Section 3.3, we consider the following regression equation:

$$\log y_{j,t} - \log y_{j,t-1} = \beta \widehat{X}_{cz(j),t} + \gamma' K_{j,t} + \epsilon_{j,t}, \tag{12}$$

[27] See https://www.glassdoor.com/blog/top-20-employee-benefits-perks-for-2017/.

where j is the firm; y is its outcome on a measure such as value added or wage bill; $cz(j)$ is its commuting zone; $\widehat{X}_{cz,t}$ denotes the *growth* in the employment share by foreign-owned firms in that commuting zone; and $K_{j,t}$ is a vector of controls discussed below. The parameter of interest is β, which is the *indirect effect*.

Identifying β is challenging for at least two reasons. First, there is a classic selection issue with the allocation of foreign multinational activity across locations. Foreign firms may choose to hire in regions in which wages are already set to grow. For example, the foreign firm may be aware of new regional investments in production infrastructure or education and increase hiring in this region to benefit from the infrastructure or workforce improvements. Then, a naive regression of earnings growth on employment growth at foreign firms would overstate the impact of foreign firm activity. Conversely, foreign firms may choose to hire in regions in which the local economy is already set to decline. For example, the foreign firm may be aware that wages or intermediate goods prices are set to decline in this region, possibly because a large existing employer plans to lay off its workforce, so the foreign firm may increase activity to take advantage of falling prices. This case is further confounded by the importance of local tax incentives, which are estimated to be large in the U.S. and may be targeted especially toward attracting foreign firms to declining regions.[28] Then, a naive regression of earnings growth on employment growth at foreign-owned firms would understate the impact of foreign firm activity.

Second, we may be mismeasuring growth in the employment share of foreign firms in the commuting zone, $\widehat{X}_{cz,t}$. As discussed in Section 2, we expect there to be some measurement error in the linkages between the parent and its subsidiaries and how these change over time.

To overcome these identification challenges, we adapt the identification strategy common in the literature about the effects of immigration on non-immigrants in the same region (Card, 2001). This literature uses the fact that immigrants cluster into regions in the U.S. based on country of origin. To adapt this instrument to identify the effects of foreign-owned firm activity on workers, we first notice that employment at foreign-owned firms tends to be clustered by region and country of origin (see Figure 1). For example, German-owned firms disproportionately employ workers in South Carolina in 2010 if they do so in 2005. This is analogous to the clustering of immigrants into regions.

We construct the instrument as the predicted change in employment at, for example, German-owned firms in South Carolina between 2009 and 2010 using only information about (i) the share of workers at German-owned firms in South Carolina in 2005 and (ii) the change in *aggregate* employment by German-owned firms in *any other* region in the U.S. between 2009 and 2010. Since this instrument is not formed using information about the change in

[28] See the discussion by Greenstone et al. (2010). Relatedly, Criscuolo, Martin, Overman, and Van Reenen (2019) find that regional investment subsidies are negatively selected in the U.K. such that naive regression estimates of their effects are severely downward biased.

employment by German-owned firms in South Carolina between 2009 and 2010, it does not depend directly on changes in South Carolina's business climate between 2009 and 2010. In other words, German firms' aggregate foreign employment growth (net of employment growth in South Carolina) in 2010 is *plausibly exogenous* of South Carolina's local unobservable shocks in 2010. In particular, it does not depend directly on infrastructure investments, improved educational opportunities, or changes in the generosity of tax incentives in South Carolina in 2010, so it does not depend directly on the confounding factors discussed above.

To formalize the approach, relative foreign-owned firm employment growth in the commuting zone, $\widehat{X}_{cz,t}$, is defined by

$$\widehat{X}_{cz,t} \equiv \frac{L^F_{cz,t} - L^F_{cz,t-1}}{L^F_{cz,t-1} + L^D_{cz,t-1}}, \tag{13}$$

where $L^F_{cz,t}$ and $L^D_{cz,t}$ are the number of employees at foreign- and domestic-owned firms in commuting zone cz and year t, respectively. The parameter of interest is the effect of a change in the regional share of employment at foreign-owned firms, $\widehat{X}_{cz,t}$, on the change in an outcome, such as the earnings growth of a worker at a domestic firm in the region.

To form the instrument, we use the tax data on the firm's country of foreign ownership to construct the share $S^o_{cz,t}$ of all employment in commuting zone cz at firms whose owners are located in origin country o, defined by

$$S^o_{cz,t} \equiv \frac{L^{F_o}_{cz,t}}{L^F_{cz,t} + L^D_{cz,t}}. \tag{14}$$

Analogous to Card (2001) and the subsequent immigration literature, we then construct the instrumental variable $\widehat{Z}_{cz,t}$ as

$$\widehat{Z}_{cz,t} = \sum_o \frac{\sum_{cz' \neq cz}(L^{F_o}_{cz',t} - L^{F_o}_{cz',t-1})}{\sum_{cz'} L^{F_o}_{cz',t-5}} S^o_{cz,t-5}. \tag{15}$$

This variable is interpreted as the prediction of $\widehat{X}_{cz,t}$, formed only from the share of employment by firms from country o in cz dated at $t-5$ and the change in aggregate employment by o in the U.S. from $t-1$ to t. Note that we modify the approach from the immigration literature slightly by leaving out own-commuting-zone employment when constructing the aggregate change from $t-1$ to t, which helps to rule out confounding factors.[29] The denominator is the total number of FTE workers in the country of origin 5 years ago, which ensures that the aggregate change is measured relative to levels dated far before contemporaneous shocks. Because $\widehat{Z}_{cz,t}$ is not a function of cz-specific changes

[29]We also consider leaving out nearby commuting zones in a robustness check (see Section 5.3).

between $t-1$ and t, it should satisfy that $\widehat{Z}_{cz,t}$ and the unexplained component of cz growth are orthogonal (conditional on observed determinants of growth $K_{j,t}$). However, we see four possible threats to identification as well as a threat to drawing inference on our estimates.

First, the instrument includes the past share of employment at foreign-owned firms from various origin countries, as well as the contemporaneous change in the employment at such firms in other regions. This raises the concern that there may be regional shocks that are correlated with our instrument. For example, regions near the Canadian border may also be affected by trade shocks originating in Canada that are correlated with the instrument. To deal with this concern, we include Census Division-year fixed effects in the regressions, which absorb all contemporaneous effects at the regional level.

Second, industry shocks may be correlated with the instrument. For example, German- or Japanese-owned firms may be more likely to be in the car industry and select commuting zones that are also abundant with other car industry firms. To deal with this concern, we also include fine industry-year fixed effects based on the 3-digit NAICS code (6-digit NAICS in a robustness check discussed in Appendix A.8) to absorb any contemporaneous nationwide growth trends by industry.

Third, foreign investment growth may be disproportionately concentrated in urban regions (see, e.g., Bakker 2020). To ensure that urban concentration does not confound the foreign shocks, we control for various measures of urban concentration, including log population size, log population density, an indicator for spatial overlap with a micropolitan statistical area, and an indicator for overlap with a metropolitan statistical area. We measure these in the pre-period to avoid controlling out the effects of interest.

Fourth, Borusyak, Hull, and Jaravel (2020) recently showed that, under their assumptions, instruments with a shift-share structure may be biased if they do not control for the sum of regional exposure shares by year. To address this, we always control for $L^D_{cz,t-5}/(L^F_{cz,t-5}+L^D_{cz,t-5})$ in the indirect effects regressions.[30]

Lastly, while it is plausible that the aggregate foreign employment growth of a country of origin (leaving out employment growth in a commuting zone) is orthogonal to local growth shocks in a particular commuting zone, this does not imply that the regression residuals are independent across nearby commuting zones. Spatially dependent residuals would not

[30] See their discussion of the "incomplete shares" problem. They also suggest interacting the domestic employment shares with time periods in order to allow for more flexible domestic shock specifications, which amounts to including more than a dozen additional linear controls in our regressions. Of course, we already allow for extremely flexible domestic shock specifications by including fine industry-year fixed effects and Census-division-year fixed effects. If we fully interact the domestic shares with years to allow even more flexibility, we find stronger indirect effects than in our baseline estimates, but the standard errors become much less precise. In Appendix Table A4, we provide a robustness check in which we interact the domestic shares with indicators for groups of years (e.g., the financial crisis of 2007-2009), where grouping the years serves as a parsimonious way to allow for additional flexibility in the domestic shocks, finding that the estimates become somewhat larger but are not statistically significantly different.

bias the regression coefficient estimate but would tend to downward bias standard errors in the regression, leading to overrejection of the null hypothesis (Adao, Kolesár, and Morales, 2019; Borusyak et al., 2020).[31] In order to be conservative when drawing inference, we follow Borusyak et al. (2020) in transforming the regression into one that is clustered at the country-of-origin-year level. However, as discussed by Borusyak et al. (2020), their method does not incorporate that the instrument leaves out own-commuting-zone employment growth. They argue that the standard errors are still approximately valid for leave-one-out point estimates. As an alternative that accounts for the leave-one-out nature of our instrument, we also provide traditional standard errors clustered at the commuting-zone-year level.

To summarize, in the baseline specification, we protect against potential confounders by including in the control vector $K_{j,t}$ industry-year indicators, Census-division-year indicators, measures of urban concentration, and the sum of commuting zone exposure shares, then report standard errors clustered at either the country-of-origin-year or commuting-zone-year level. In Appendix A.8, we demonstrate that the results are not sensitive to adding control variables.

5.2 Estimates of Indirect Effects on Local Labor Markets

We next discuss our baseline estimates of indirect effects. The instrument and endogenous variable are constructed from information on both foreign and domestic firms, while the sample in the regression includes only continuing domestic firms.[32] The outcomes of interest are value added, employment, wage bill, and earnings of continuing workers at domestic firms, and the sample size may vary across outcomes. (For example, value added can be negative, in which case log value added is not defined.) All observations are weighted by the number of FTE workers in $t-1$. The control variables were discussed in the previous subsection.

The full sample results are presented in the first column of Table 1. The first-stage coefficient is 0.56. The F-statistic is above 230 when clustering by commuting-zone-year and above 40 when clustering conservatively by country-of-origin-year. Thus, lagged shares of foreign employment by country of origin in a commuting zone interacted with that country's aggregate employment growth provides an economically and statistically significant predictor

[31] Adao et al. (2019) summarize the overrejection issue as follows: "Whenever two regions have similar exposure shares, they will not only have similar exposure to the aggregate shocks, but will also tend to have similar values of the regression residuals. While traditional inference methods allow for some forms of dependence between the residuals, such as spatial dependence within a state, they do not directly address the possible dependence between residuals generated by unobserved shift-share components.... [T]raditional inference methods underestimate the variance of the OLS estimator of β, creating the overrejection problem."

[32] The outcome sample includes both domestic multinationals and domestic non-multinationals. We find that results are similar when restricting the sample to domestic non-multinationals in a robustness check (see Section 5.3).

of that country's employment growth in the commuting zone. Using the instrument, we estimate that a 1 percentage point increase in the share of employment at foreign firms in the commuting zone increases the value added, employment, and wage bill at domestic firms by 0.96 percent, 0.53 percent, and 0.63 percent, respectively.[33] These estimates are statistically significant at the 0.01 significance level even when using conservative standard errors clustered at the country-of-origin-year level. Appendix Table A5 compares these estimates to what we would obtain using OLS, with and without our rich set of controls. We find that the OLS estimates are about half as large as the estimates using our instrumental variable. As we discussed in the previous subsection, one reason for OLS estimates to be smaller is measurement error in $\widehat{X}_{cz,t}$; another reason is the selection of foreign investment into declining regions induced, for example, by tax incentives or declining prices.

We also examine indirect effects on earnings at the worker-level. To do so, we perform a regression for continuing workers in the same domestic firm and commuting zone. We use a within-worker differenced specification to remove both worker fixed effects and firm fixed effects. The regression controls are the same as above, except for individuals instead of firms as the observations, and a polynomial in age is included to control for heterogeneous age profiles in earning growth. The results are presented in Panel D of Table 1 for about 370 million worker-year observations. The full sample estimate indicates a positive and statistically significant effect on the average worker's earning growth of about 0.15. This is greater than the estimate of 0.10 that one would obtain using the difference between log wage bill and log employment effects in Panels B-C, highlighting the importance of controlling for worker composition in order to understand the earning growth effects of foreign investment.

Next, we consider heterogeneity in the effects across firm types using the same empirical specification but applied to various subsamples. Columns 2-4 of Table 1 explore heterogeneity in the indirect estimates for three size groups, using the number of FTE workers measured at $t-1$. Columns 5-6 consider heterogeneity in the effect on tradable versus non-tradables industries, using the classifications of Mian and Sufi (2014). We then repeat the regression in (12) for each of these groups of firms. We find that the effects are much larger among large firms and firms in the tradable sector. We estimate that a 1 percentage point increase in the share of employment at foreign firms in the commuting zone increases value added by 2.7 percent at firms with at least 100 workers and by 3.4 percent at firms in the tradable sector. By contrast, the point estimate is small and insignificant for firms with fewer than 10 workers and is smaller yet still statistically significant in the non-tradables sector. The patterns are similar for the FTE employment, wage bill, and earnings of continuing workers.[34]

[33] Note that the indirect effect estimates are semi-elasticities. In Section 6, we convert these estimates to dollars or jobs generated at domestic firms in response to one additional job created at a foreign firm.

[34] Iacovone, Javorcik, Keller, and Tybout (2015) find qualitatively similar differences of the effects of FDI growth on domestic firms by firm size. They find negative effects from Walmart's entry into Mexico on small

Table 1: Indirect Effect Estimates: Main Results

	Full Sample	**By Firm Size**			**By Sector**	
		Size 1-9	**Size 10-99**	**Size 100+**	**Tradables**	**Non-tradables**
Panel A.	**Outcome: Log Change in Value Added**					
Second-Stage Coefficient	**0.96**	**0.12**	**0.54**	**2.66**	**3.38**	**0.50**
(Std. Error Clustered by Commuting Zone)	(0.30)	(0.10)	(0.18)	(1.14)	(1.56)	(0.23)
(Std. Error Clustered by Country of Origin)	(0.51)	(0.08)	(0.20)	(1.64)	(3.10)	(0.23)
First-Stage Coefficient	0.56	0.59	0.53	0.49	0.53	0.48
(F-statistic Clustered by Commuting Zone)	(232)	(361)	(241)	(112)	(128)	(143)
(F-statistic Clustered by Country of Origin)	(42)	(40)	(52)	(65)	(46)	(51)
Number of Firms by Commuting Zones (Millions)	41.8	34.9	6.5	0.5	6.0	6.0
Number of Workers (Millions, measured at $t-1$)	416.8	96.2	158.5	162.2	98.3	63.3
Panel B.	**Outcome: Log Change in Employment**					
Second-Stage Coefficient	**0.53**	**0.02**	**0.40**	**1.55**	**1.22**	**0.72**
(Std. Error Clustered by Commuting Zone)	(0.14)	(0.08)	(0.16)	(0.52)	(0.43)	(0.25)
(Std. Error Clustered by Country of Origin)	(0.18)	(0.07)	(0.17)	(0.54)	(0.43)	(0.26)
First-Stage Coefficient	0.56	0.59	0.53	0.50	0.53	0.48
(F-statistic Clustered by Commuting Zone)	(235)	(364)	(246)	(119)	(130)	(143)
(F-statistic Clustered by Country of Origin)	(44)	(39)	(52)	(66)	(49)	(53)
Number of Firms by Commuting Zones (Millions)	46.0	38.3	7.1	0.5	6.4	6.4
Number of Workers (Millions, measured at $t-1$)	477.3	105.1	175.8	196.5	107.3	71.1
Panel C.	**Outcome: Log Change in Wage Bill**					
Second-Stage Coefficient	**0.63**	**0.00**	**0.41**	**1.62**	**1.42**	**1.19**
(Std. Error Clustered by Commuting Zone)	(0.17)	(0.10)	(0.18)	(0.53)	(0.47)	(0.35)
(Std. Error Clustered by Country of Origin)	(0.22)	(0.10)	(0.19)	(0.56)	(0.50)	(0.36)
First-Stage Coefficient	0.56	0.59	0.53	0.50	0.53	0.48
(F-statistic Clustered by Commuting Zone)	(235)	(364)	(246)	(119)	(130)	(143)
(F-statistic Clustered by Country of Origin)	(44)	(39)	(52)	(66)	(49)	(53)
Number of Firms by Commuting Zones (Millions)	46.0	38.3	7.1	0.5	6.4	6.4
Number of Workers (Millions, measured at $t-1$)	477.3	105.1	175.8	196.5	107.3	71.1
Panel D.	**Outcome: Log Change in Earnings of Continuing Workers**					
Second-Stage Coefficient	**0.15**	**0.01**	**0.06**	**0.40**	**0.39**	**0.19**
(Std. Error Clustered by Commuting Zone)	(0.07)	(0.05)	(0.07)	(0.14)	(0.16)	(0.09)
(Std. Error Clustered by Country of Origin)	(0.08)	(0.07)	(0.08)	(0.15)	(0.17)	(0.09)
First-Stage Coefficient	0.56	0.59	0.54	0.50	0.53	0.49
(F-statistic Clustered by Commuting Zone)	(239)	(367)	(249)	(123)	(134)	(149)
(F-statistic Clustered by Country of Origin)	(44)	(39)	(52)	(66)	(48)	(53)
Number of Firms by Commuting Zones (Millions)	44.6	37.0	7.1	0.5	6.3	6.2
Number of Workers (Millions, measured at $t-1$)	369.6	83.4	130.9	155.3	87.2	54.4

Notes: The outcome sample only includes continuing domestic firms. Observations are weighted by lagged firm size. Controls are industry-year indicators, Census-division-year indicators, measures of urban concentration, and the sum of commuting zone exposure shares.

Lastly, to investigate inequality in the worker-level earnings effects, we split the sample into equally-sized quintile bins by ranking lagged earnings within the commuting-zone-year. In columns 2-6 of Table 2, we examine earnings growth effects for continuing workers at

Mexican suppliers of retailers and positive effects on large suppliers.

Table 2: Indirect Effect Estimates: Results by Worker Earnings Quintile

	Full Sample	By Earnings Quintile Group: Quintile 1	Quintile 2	Quintile 3	Quintile 4	Quintile 5
	Outcome: Log Change in Earnings of Continuing Workers					
Second-Stage Coefficient	**0.15**	**0.06**	**0.04**	**0.08**	**0.27**	**0.32**
(Std. Error Clustered by Commuting Zone)	(0.07)	(0.12)	(0.09)	(0.08)	(0.09)	(0.12)
(Std. Error Clustered by Country of Origin)	(0.08)	(0.13)	(0.12)	(0.09)	(0.10)	(0.12)
First-Stage Coefficient	0.56	0.55	0.56	0.56	0.56	0.55
(F-statistic Clustered by Commuting Zone)	(239)	(235)	(237)	(238)	(238)	(238)
(F-statistic Clustered by Country of Origin)	(44)	(50)	(47)	(47)	(46)	(47)
Number of Firms by Commuting Zones (Millions)	44.6	20.1	19.6	18.8	17.0	16.1
Number of Workers (Millions, measured at $t-1$)	369.6	73.9	73.9	73.9	73.9	73.9

Notes: The sample includes only workers employed by the same domestic firm in the same commuting zone during t and $t-1$. The sample only includes continuing workers at domestic firms. We divide workers into five earnings groups within each commuting-zone-year based on the ordering of their lagged earnings. Controls are industry-year indicators, Census-division-year indicators, measures of urban concentration, and the sum of commuting zone exposure shares.

different lagged earning quintile bins. For the lowest three quintile bins, we find positive but statistically insignificant estimates. For the top two quintile bins, we find statistically significant estimates of about 0.3. This indicates that a 1 percentage point increase in the share of employment at foreign firms in the commuting zone results in 0.3 percent wage growth for high-paid continuing workers at domestic firms in the commuting zone, while low-paid workers experience little to no wage growth. This implies indirect effects primarily benefit high-skilled workers at domestic firms, as predicted by our model (see Section 3).

5.3 Robustness of the Indirect Effect Estimates

In Appendix A.8, we provide numerous robustness checks to address potential concerns with the research design, which we briefly summarize here. In a placebo test in which domestic firms' outcomes are measured in the pre-period, the estimated effects become small in magnitude and statistically insignificant for all of the outcomes, consistent with our identifying assumption. Next, a potential concern with shift-share instruments is that the second-stage coefficient may conflate the effects of contemporaneous and past shocks if the shocks have delayed impacts (Jaeger, Ruist, and Stuhler, 2018). We check that our estimates are nearly identical when controlling for the lagged shocks, implying that our results are not confounded by delayed impacts. Furthermore, our findings are robust to leaving out any commuting zone within a 300-mile radius of the worker's residence when constructing the shocks, indicating that the estimates are not confounded by the possibility of workers responding to shocks in nearby regions. Excluding all 52 countries that Hines (2010) considers tax havens, we find similar estimates, indicating that misclassification of some domestic firms as foreign for tax avoidance purposes does not bias our findings. Another possible threat to identification is that aggregate employment growth from a country of origin may lower transportation costs for U.S. exports to that country. Since most U.S. exports are carried out by multinationals (Bernard, Jensen, and Schott, 2005), we check if the estimates conflate foreign demand effects with foreign employment effects by dropping domestic multinationals from the outcome sample, finding that the estimates are unaffected. Finally, to incorporate entry and exit into the outcome measures, we consider the transformation of Davis, Haltiwanger, and Schuh (1998). The estimated effects become somewhat stronger, which ameliorates any concern that our main effects for continuing firms arise from survival bias.

5.4 Understanding the Mechanisms behind the Indirect Effects

We conclude this section by discussing a number of mechanisms that could explain the positive indirect effects estimates. In our model in Section 3, positive indirect effects arise from knowledge spillovers from foreign to domestic firms. We first note that knowledge spillovers could come in the form of technology or improved management practices. Bloom et al. (2019) find evidence for local spillovers in management practices associated with large plant openings using the "Million Dollar Plants" research design. In fact, most million dollar plants in their study belong to multinational corporations.

Outside the scope of our model, increased competitive pressure may lead to higher efficiency at domestic firms (see Bloom, Draca, and Van Reenen 2015). However, competitive pressure would also predict that these firms become smaller in the short-run, contrary to our results. Yet another channel for positive indirect effects on local domestic firms is an increase in consumer demand for non-tradables (see Moretti 2010). While we do find sizable effects in this sector, the effects are even greater in the tradable sector—suggesting that consumer demand cannot be the only channel behind the indirect effects.

Another potential mechanism through which indirect effects may arise is the firms' input-output network (see Aitken and Harrison 1999 and Javorcik 2004). Increased foreign investment may result in cheaper intermediate inputs supplied to domestic firms or greater local demand for the output of domestic firms. Either would likely result in greater output and employment at domestic-owned firms, so input/output spillovers can be thought of as an alternative interpretation of the productivity spillovers in our model. Javorcik (2004) investigated spillovers at the national level in Lithuania and found primarily positive effects from foreign investment on upstream domestic firms. Similarly, Alfaro-Urena et al. (2019b) find positive productivity effects for domestic firms selling to multinational firms in Costa Rica.[35]

[35] In an earlier draft of this paper, we provided estimates of upstream and downstream effects when using industry-level input/output tables to measure exposure to upstream and downstream foreign investment

6 Local and Aggregate Implications

In this section, we use our estimates from Sections 4 and 5 to take a look at the local and aggregate implications of foreign multinationals. We emphasize that the numbers calculated below are not meant to summarize the overall welfare effect of foreign multinationals. We abstract, for example, from any worker-firm-specific preference heterogeneity in the calculations below. The calculations below are based on aggregate outcomes in 2015.

6.1 Aggregate Direct Effects

We start by conducting the following thought experiment: Suppose one replaces all foreign multinationals with domestic firms—each equipped with the average productivity of domestic firms. How much would this lower the aggregate wages in the U.S.? We abstract away from any indirect effects (e.g., local spillovers) or worker-firm interactions.[36] In Section 4, we estimate an average foreign wage premium of 7 percent—after removing the effect of worker skill differentials from the wage differential between foreign and domestic firms. The theory suggests that this wage premium arises because of the larger productivity of foreign firms. Given an aggregate wage bill at foreign multinationals in the U.S. of 515 billion USD, this suggests an aggregate national wage premium due to foreign multinationals in the ballpark of 36 billion USD annually.[37] These figures suggest large aggregate gains for workers in the U.S. because of foreign multinationals. Indeed, 36 billion USD exceeds the aggregate subsidies of 4.6 billion USD paid to foreign firms per year.

6.2 Local Effects of a New Foreign Plant

Beyond aggregate wage effects, policymakers are often confronted with weighing the local economic benefits of a foreign firm against subsidy costs. To be concrete, consider the establishment or expansion of a foreign firm that would create 1,000 new jobs in a commuting zone. Unlike in the previous subsection, we do not compare this expansion to a domestic firm expansion of similar size. The reason is that here we are interested in the direct as well as the local indirect effects, and our identification strategy delivers the indirect effects of foreign firms but not of domestic firms. Hence, the thought experiment is having a new

shocks. However, due to the absence of firm-to-firm transactions data in U.S. tax records, it is not possible to precisely measure upstream and downstream exposure at the firm-level, and our instrument lacked the statistical power to distinguish between these channels. We hope that firm-to-firm transactions data will one day be available for the U.S. so that this analysis can be performed.

[36] By comparing one commuting zone with another above, we estimate the local indirect effects of foreign firms but not the national indirect effects, which are differenced out.

[37] We calculate the aggregate wage bill at foreign multinationals from the average wage of a full-time employee at foreign-owned firms (Table A1) and the number of workers at foreign multinationals from the BEA (6.8 million). We use per-worker estimates from tax data, but we use BEA aggregate estimates because it is not possible to link all workers to firms in the tax data, as discussed in Section 2.

foreign plant with 1,000 jobs compared to not having a new plant. Below, we describe some of the expected direct and indirect local effects. We focus on a commuting zone with an initial employment share of 94 percent at domestic firms, which corresponds to the national average. The benefits estimated in this subsection are calculated from the perspective of a local policymaker, while the previous subsection on aggregate direct effects is from the perspective of a national policymaker. While a local policymaker considers it valuable to steal business from another location, a national policymaker would discount the benefits of cross-location business stealing. See the discussion by Glaeser and Gottlieb (2009).

Wage gains for domestic incumbents. Since 87 percent of workers who are hired by foreign multinationals from domestic-owned firms were previously employed in the same commuting zone, our calculations assume that around 870 of the 1,000 new positions will be filled by domestic incumbents. From our foreign firm premium estimate, direct wage gains for domestic incumbents hired by the foreign firm sum to 4.6 million USD.[38] Wage gains for domestic incumbents also include those that arise indirectly at domestic firms. Recall that we estimate a wage increase of 0.15 percent for workers at domestic firms, due to a 1 percentage point increase of the share of employment at foreign firms (see Table 2). The average earning of a full-time employed worker at a domestic firm is 62,600 USD. Combining these figures suggests an indirect wage effect of 8.8 million USD for domestic incumbents who are not hired by the foreign firm.[39] In total, we find a 13.4 million USD wage gain for domestic incumbents due to 1,000 hires by a foreign firm, or 13,400 USD per created job, of which two-thirds is from the indirect effects.[40]

Increase in local economic activity. Beyond affecting the wages for incumbents, foreign multinationals also affect the overall size of economic activity in a location. While the theory suggests that the indirect effects on output at domestic-owned firms can be positive or negative, the empirical analysis in Section 5 suggests that the local indirect effects are positive on average. We calculate that 1,000 positions at a foreign-owned plant on average raise the value added in the commuting zone by 359 million USD per year.[41] Furthermore, employment increases by around 1,500 positions (i.e., an indirect effect of 500 more jobs at

[38]Specifically, 870 workers × 75,700 USD per worker × 7 percent = 4.6 million USD.

[39]Let ζ denote the commuting zone size. 94 percent of ζ workers experience a $0.15 \times \frac{1{,}000}{\zeta} \times 62{,}600$ USD wage gain, resulting in an indirect gain of 8.8 million USD for this group of workers.

[40]Note that, on a per-job basis, the results are independent of the magnitude of the increase in employment at foreign owned firms and independent of commuting zone size. The effects get slightly larger with a smaller fraction of initial employment at foreign-owned firms in the commuting zone.

[41]The value added per worker at a foreign multinational is 220,100 USD and 154,300 USD at a domestic firm on average. In addition to a direct increase in value added in the commuting zone by 220 million USD, the estimates in Table 1 suggest an indirect increase in value added by 139.2 million USD (calculated as $\frac{1{,}000}{\zeta} \times 0.96 \times 0.94 \times \zeta \times$ 154,300 USD).

domestic firms), and the total wage bill increases by 112.8 million USD on average.[42] Our estimate of a total local job multiplier of about 1.50 (0.50 indirect jobs for each 1 job created) is at the lower end of estimates in the urban economics literature, which typically range from 1.5 to 2.5 (see the review by Bartik and Sotherland 2019). While the literature lacks a directly comparable estimate on the job multiplier for foreign multinationals, Moretti (2010) finds that, for each job created in the tradable sector, 1.6 jobs are created in the non-tradable sector, for a total job multiplier of 2.6.

Comparison to local subsidies. As discussed above, our estimates do not shed light on the national indirect effects of foreign firms, but they do shed light on the local indirect effects. These calculations are still policy-relevant, as local governments actively engage in subsidy competition to attract firms (see Gaubert 2018 and Ossa 2017). Extending data collected by the policy group Good Jobs First, Slattery (2020) analyzes 387 large subsidy deals given by state and local governments in the U.S. In these data, firms promise to create 1,400 direct jobs and receive a subsidy worth 150 million USD on average, so these mega-deals are a natural comparison for our hypothetical 1,000 job plant. About a quarter of these large subsidy deals go to foreign multinationals and the median subsidy per direct job given to a foreign parent is 100,000 USD.[43] Our estimate of 13,400 USD annual wage benefits to domestic incumbents is a conservative estimate of total benefits, as it omits other non-wage benefits to the commuting zone (e.g., increased tax revenues, increased variety of employment options). At a discount rate of about 0.13, the average wage benefits per position at a foreign firm equal the typical subsidy payment. At a discount rate of 0.10, the net present value of the average wage gain exceeds the typical subsidy by 34,000 USD per job. Since foreign multinationals are mobile in their location choices for large plant openings or expansions, it is intuitive that in the bargaining with local authorities over mega-deals, they typically extract a large fraction of the overall local benefits via subsidy payments.

[42] The estimates in Panel B of Table 1 suggest an indirect increase in employment of about 500 workers (calculated as $\frac{1{,}000}{\zeta} \times 0.53 \times 0.94 \times \zeta$). If the foreign employment share is zero, the predicted indirect increase rises to 530 workers. The foreign plant would lead, on average, to a direct increase in the wage bill at foreign-owned firms of 75.7 million USD. Using the estimates in Panel C of Table 1, we compute an indirect increase in the wage bill at domestic owned firms of 37.1 million USD (calculated as $\frac{1{,}000}{\zeta} \times 0.63 \times 0.94 \times \zeta \times$ 62,600 USD). The increase in the total wage bill is substantially larger than the increase in the wage premium for incumbents calculated in Section 6.2 above, as it includes wages paid to individuals that were previously working outside the commuting zone or were non-employed.

[43] The median subsidy given to a U.S. parent is 60,000 USD. We are grateful to Cailin Slattery for providing these statistics.

7 Conclusions

In this paper, we use employer-employee panel data from 1999 to 2017 to conduct a comprehensive analysis of the effects of foreign multinationals in the U.S. We find that these firms pay a wage premium of about 7 percent on average, meaning that the same worker earns 7 percent more at a foreign-owned firm. The wage premium is larger for higher-skilled workers and absent for the lowest decile of worker skill. Our theory rationalizes these findings with a (skill-biased) productivity advantage of foreign firms. Empirically, we document that this foreign firm premium is correlated with the GDP per capita of the origin country. Furthermore, on average, the firm premium is about the same for domestic multinational firms, suggesting that the multinational status itself is associated with higher wages for the same worker. Quantitatively, the wage premium paid by foreign multinationals is quite large in the aggregate—accounting for 36 billion USD annually in wages (which is about 0.6 percent of the entire private sector wage bill). Though we did not find that controlling for measures of local and national employment would substantially reduce the multinational wage premium, we do not observe a multinational firm's global employment size. In future work, it would be interesting to evaluate how much of the multinational wage premium arises from economies of scale associated with its global employment size.

In terms of policy implications, our estimates highlight sizable benefits of trade and investment policies that make it attractive for foreign firms to invest in the U.S. Furthermore, our estimates imply incentives for local policymakers to compete for investments by foreign multinationals, since, in addition to direct wage benefits, we find positive and sizable local indirect effects on domestic firms and their workers—in particular, the higher-earning ones. We note that while it is rational for local policymakers to compete for foreign multinational investments with subsidies, this does not imply that such subsidies are beneficial from a national welfare perspective. Our calculations suggest that the subsidies given to foreign multinationals for large plant investment or expansions account for a sizable fraction of the net present value of the wage benefits for incumbent workers. In other words, foreign multinationals are able to extract a sizable fraction of the surplus from such investments in the bargaining with local governments over mega-deals.

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A Online Appendix

A.1 Data Sources and Variable Definitions

Worker data. Worker data are constructed from annual Form W-2 tax filings over the years 1999-2017.

- Worker identifier: The worker is identified by the taxpayer identification number (TIN), which is unique and allows us to follow the same worker over time and across firms. In our data, the TIN is masked to protect confidentiality.

- Employer: Form W-2 is filed by the firm on behalf of the worker and includes that firm's masked employer identification number (EIN), which we use to link workers to their employers. In the event that multiple EINs file Form W-2 for the same TIN in year t, we define the EIN with the greatest earnings as the employer in year t, as is standard in the literature on firm-worker panel data.

- Earnings: Reported on Form W-2, box 1, earnings are defined as all remuneration for labor services deemed taxable by the IRS, including wages and salaries, bonuses, tips, and exercised stock options. Following Lamadon et al. (2020), the analysis sample focuses on workers with earnings above the full-time equivalence (FTE) threshold, approximated by the minimum wage, which equates to 15,000 USD in 2015. Note that we observe annual earnings, but since workers do not report hours worked in tax records, it is not possible to construct a measure of the hourly wage. To protect against outliers, we winsorize both log earnings and changes in log earnings from above and below at the one-half percent level.

- Location: Form W-2 reports the residential ZIP code of the worker. We define the location as the commuting zone associated with this ZIP code using the year 2000 commuting zone definitions from the Bureau of Labor Statistics. In the event that the ZIP code is missing or invalid in year t but not in year s with $|t - s| \leq 2$, and the worker receives a W-2 from the same EIN in t and s, we impute it in t using the value from s.

- Age: We obtain year of birth from SSA birth records. Following Lamadon et al. (2020), the analysis sample focuses on workers between age 25 and 60.

Firm data. Firm data are constructed from annual business tax returns over the years 1999-2017. The source tax forms are Form 1120 (C-corporations), Form 1120-S (S-corporations), and Form 1065 (Partnerships). We improve the data by imputing industry codes from other tax forms when missing, correcting value added for the particular industries that partially deduct labor costs, and using subsidiary links to associate foreign ownership with each subsidiary instead of only the parent corporation. Exhaustive variable definition and improvement steps are as follows:

- Firm identifier: A unique firm in the business tax filings is defined by the employer identification number (EIN). The EIN is the level at which companies file their tax returns with the IRS, so it reflects a distinct business unit for tax and accounting purposes. The EIN is often, but not always, the parent corporation in a multi-establishment firm. See Song et al. (2018), who also define the firm as the EIN, for further discussion of differences between EINs and establishments. In our data, the EIN is masked to protect confidentiality.

- Foreign ownership: We define an EIN as foreign owned in year t if it files Form 5472 in year t. Form 5472 is the "Information Return of a 25% Foreign-Owned U.S. Corporation or a Foreign Corporation Engaged in a U.S. Trade or Business." The country of foreign ownership is also reported on Form 5472. Note that S-corporations were restricted by law to only be owned by U.S. citizens during our time frame. Note that even a domestic-owned firm could be in the hands of many small foreign owners, particularly, when the company is publicly listed. We do not have hard data on this, but we think these cases are likely to be rare and not necessarily associated with the same effects. In the event that the employer fails to file Form 5472 in year t but files as foreign owned with ownership country c in one of $(t-2, t-1)$ as well as one of $(t+1, t+2)$, we improve the data by imputing foreign ownership in year t as c.[44] Since we do not observe the previous year for the initial year of the sample, we cannot carry out the same imputation and exclude the initial year from the estimation.

- Multinational: We define an EIN as a multinational in year t if it reports a non-zero foreign tax credit on Schedule J, Part I, line 5a of Form 1120 or Form 1118, Schedule B, Part III, line 6 of Form 1118 for a C-corporation in year t, or if it reports a positive Total Foreign Taxes Amount on Schedule K, line 16l of Form 1065 for a partnership in year t, while S-corporations are restricted by law from carrying out foreign business.

- Subsidiary: As emphasized by Yagan (2019), many workers cannot be linked to a corporate tax filing, often because the employer is not required to file (especially because

[44] An additional issue that may result in measurement error is that some firms may outsource their employee administration to third-party payroll processors whose EINs appear on the W-2 rather than the EINs of the actual employers. In this case, we would treat the payroll processor as a separate employer, rather than combining it with the firm that directly employs the workers, since we do not have a way of mapping payroll processors back to direct employers. However, as noted by Yagan (2019), only a small number of firms is likely to use the EINs of payroll processors.

the employer is a government or non-profit organization) or because the employer is a subsidiary and only the parent corporation files while the subsidiary uses its distinct EIN to issue W-2 forms. To overcome this challenge, we combine two sources of information on subsidiary linkages. The first source is Schedule K, line 3b, which provides the EIN of the parent corporation in the years in which the subsidiary is a filer, from which we learn the EIN of the parent corporation in future years in which the subsidiary is a non-filer. The second source is the Affiliations Schedule from Form 851, which defines a subsidiary as 80 percent owned by another corporation. However, we only observe a running list of parent-subsidiary relationships taken from the Affiliations Schedules through 2016, so changes over time due to extensive margin changes in subsidiary relationships may be mismeasured when using the second source. For this reason, we only utilize the second source for subsidiary linkages that are not covered by the first source (i.e., subsidiaries that are missing Schedule K filings).

- Industry: The industry of the firm in year t is reported as the 6-digit NAICS code on line 21 on Schedule K for C-corporations, line 2a Schedule B for S-corporations, and Box A for partnerships in year t. In the baseline specification, we consider the 3-digit NAICS code to be the industry, while we consider the 6-digit NAICS code in robustness checks. In the event that the NAICS code is missing in year t, we impute the NAICS code in year t-1, t-2, t+1, or t+2 (in that order). In the event that the NAICS code is missing in all such years, we attempt to impute the NAICS code from Form 5500, "Annual Return/Report of Employee Benefit Plan," as this filing sometimes includes the NAICS code even when the main business filing does not. In the data, we find that a large share of foreign-owned firms are concentrated in NAICS sector 55, "management of other companies," while very few domestic firms belong to this sector. Because sector 55 does not correspond to any particular product market, it is difficult to define its upstream or downstream industries. To avoid losing much of the sample of foreign-owned firms in the input/output network regression, we use the NAICS code of the largest subsidiary to replace a NAICS code beginning with 55 if a different NAICS is available at the largest subsidiary. Lastly, we omit the finance, insurance, and real estate (FIRE) industries throughout all analysis because of the difficulties in interpreting value added for these industries.

- Tradables and Non-tradables: Mian and Sufi (2014) provide two methods for defining the tradable industries. We say an industry is tradable if either: (A) the industry has imports plus exports equal to at least $10,000 per worker, or if total exports plus imports for the NAICS four-digit industry exceed $500M; or (B) the industry has a high level of geographic concentration (i.e., is in the highest quartile of the geographical

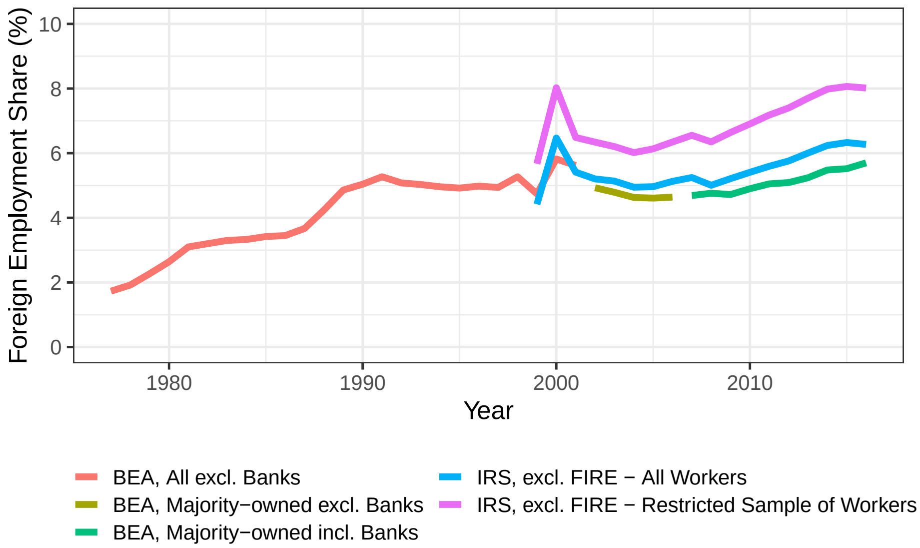
Herfindahl index constructed by Mian and Sufi 2014). We define an industry as non-tradable if it belongs to the retail sector or restaurants (corresponding to the first classification by Mian and Sufi 2014).

- Value added: We define value added as the difference between gross business receipts and the cost of goods sold (COGS). This difference is reported on line 3 for Forms 1120, 1120-S, and 1065. The IRS provides instructions to businesses on the calculation of COGS in Publication 334. To quote this publication, "Labor costs are usually an element of cost of goods sold only in a manufacturing or mining business. Small merchandisers (wholesalers, retailers, etc.) usually do not have labor costs that can properly be charged to cost of goods sold. In a manufacturing business, labor costs properly allocable to the cost of goods sold include both the direct and indirect labor used in fabricating the raw material into a finished, saleable product." Labor expenses are not included in COGS—and therefore are not subtracted out of gross business receipts when defining value added—for any business that does not engage in manufacturing or mining. Among firms that engage in manufacturing and mining, labor expenses are included for workers engaged in production ("production workers"), but not for workers who are not engaged in production ("non-production workers"). Form 1125-A is not available to us, so we do not observe the labor expense for production workers. However, we are able to recover the non-production wage and salary expenses (lines 12 plus 13 for Form 1120, lines 7 plus 8 for Form 1120S, and lines 9 plus 10 for Form 1065). We observe total wage and salary expenses from the worker data discussed below. The difference in total wage and salary expenses and non-production wage and salary expenses is production wage and salary expenses. Thus, we are able to add production wage and salary expenses into the line 3 measure for the manufacturing and mining industries (NAICS codes beginning 31, 32, 33, or 212) in order to recover value added for these industries. To protect against outliers, we winsorize changes in log value added from above and below at the three percent level.

- Location: Our analysis requires a firm's activity to be associated with *each* commuting zone in which it is active. This differs from using the address of the firm's headquarter to define its location, as the headquarter may be chosen to obtain favorable state-level tax rates rather than to represent the firm's actual location of activity, and the firm may be active in many locations. Since specific establishments of multi-establishment firms are not observable in U.S. tax data, we follow Yagan (2019) by inferring firms' commuting zone-level operations from workers' residential locations. We aggregate the number of workers and wages within the commuting zone of the worker's address on the W-2 to define the firms' local employment and wage bill. However, we do not observe value added at the firm-commuting zone level directly because it is reported only on EIN-level tax forms. To overcome this challenge, we use the share of the wage bill paid in the commuting zone of each firm to allocate value added to commuting zones. For example, if 75 percent of a firm's wage bill is paid in the first commuting zone and 25 percent in the second commuting zone, we allocate 75 percent of value added to the first and 25 percent to the second.

A.2 Descriptive Statistics

Figure A.1: Employment at Foreign-owned Firms



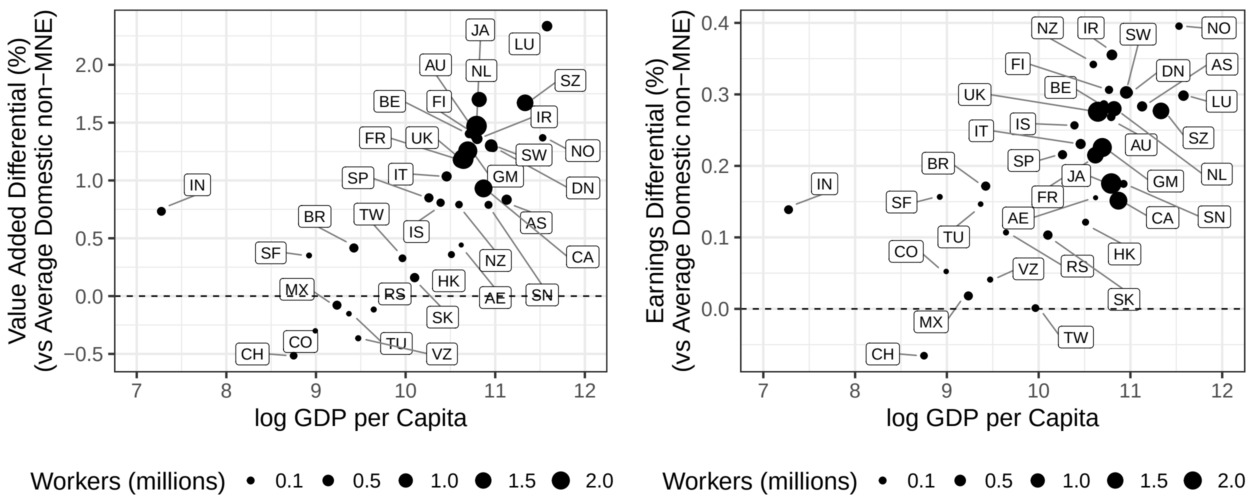
Notes: This figure displays the share of American private sector employees at foreign-owned firms between 1977 and 2017. It compares three series available from BEA to the analysis sample of firms we construct from tax data, both for all workers and for only the workers that satisfy our FTE and other restrictions. Each of the series use different sample selection rules.

Table A1: Descriptive Statistics for the Main Sample of Firms, 2015

	Domestic	Foreign
Firms in Main Sample of Firms (thousands)	2,781.1	30.3
Firm-Location Pairs in Main Sample of Firms (thousands)	4,762.9	218.7
Number of Workers at Main Sample of Firms (millions):		
All Workers:	77.1	5.2
FTE Analysis Sample:	41.3	3.6
Mean Wage at Main Sample of Firms (thousands):		
All Workers:	41.4	60.7
FTE Analysis Sample:	62.6	75.7
Value Added per Worker at Main Sample of Firms (thousands):		
All Workers:	82.7	153.1
FTE Analysis Sample:	154.3	220.1

Notes: This table displays descriptive statistics for domestic and foreign filers of Forms 1120, 1120-S, and 1065, matched to subsidiaries and W-2 forms. The set of firms is the same across all rows and has already been restricted to satisfy the sample restrictions. The analysis sample restrictions on the workers are at least FTE earnings ($15,000 per year), the firm is the worker's highest-paying W-2 in that year, the worker is prime age (25-60 years old), and the ZIP code is non-missing and valid on the highest-paying W-2 form.

Figure A.2: Descriptive Statistics by Country of Ownership

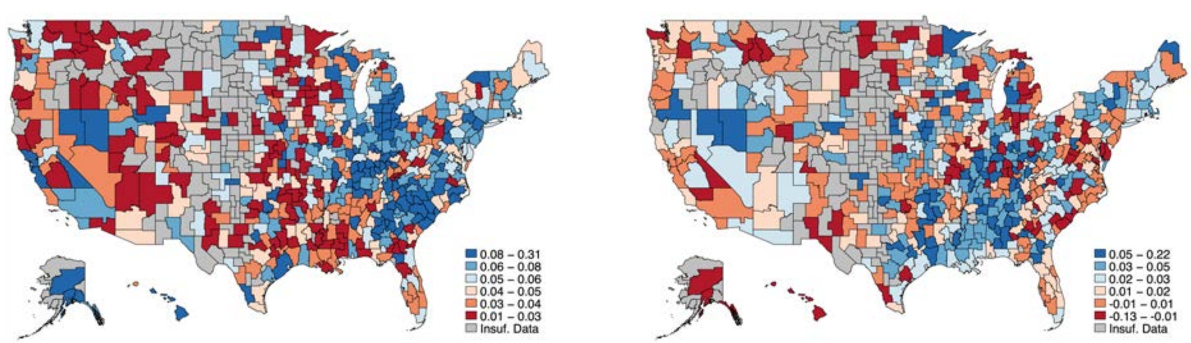


(a) Value-added differential

(b) Wage differential

Notes: This figure presents average value added and earnings during 2010-2015. The vertical axis is the difference in the average value added (subfigure a) or average earnings (subfigure b) for foreign multinationals with the countries of ownership indicated by the labels, relative to the average domestic non-multinational. We control for industry-year and commuting-zone-year fixed effects, so reported differentials in log value added and log earnings do not reflect differences due to location or industry selection.

Figure A.3: The Spatial Distribution of Employment at Foreign Firms



(a) Share of employment at foreign firms by commuting zone in 2001

(b) Change in share of employment at foreign firms by commuting zone from 2001 to 2015.

Notes: The two figures display spatial variation in employment at foreign-owned firms observed in the tax data for the workers sample of interest. In the first figure, the share of workers employed at foreign-owned firms is plotted in 2001 for each commuting zone. In the second figure, changes from 2001 to 2015 in the share of employment at foreign-owned firms are plotted by commuting zone.

A.3 Model Derivations

We now provide details on the model and prove several claims made in Section 3.

Wage setting. Recall that all firms produce the same homogeneous good whose price is normalized to one. Each firm solves the following problem:

$$\max_{w_{js},w_{ju}} \phi_j \left(w_{ju}^{\eta} \left(\frac{\bar{L}_u}{W_u} \right) + \zeta_{js} w_{js}^{\eta} \left(\frac{\bar{L}_s}{W_s} \right) \right) - w_{js}^{\eta+1} \frac{\bar{L}_s}{W_s} - w_{ju}^{\eta+1} \frac{\bar{L}_u}{W_u}. \tag{A1}$$

The first-order condition that ignores any effect of w_{js} and w_{ju} on $\frac{\bar{L}_s}{W_s}$ and $\frac{\bar{L}_u}{W_u}$ is simply equation (5).

Mean difference in log wages between foreign and domestic firms.

$$\begin{aligned} \mathbb{E}\left[\log w_{F\cdot}\right] - \mathbb{E}\left[\log w_{D\cdot}\right] &= C_F \log w_{Fs} + (1 - C_F) \log w_{Fu} - C_D \log w_{Ds} - (1 - C_D) \log w_{Du} \\ &= \log \phi_F - \log \phi_D + C_F \log \zeta_{Fs} - C_D \log \zeta_{Ds} \end{aligned} \tag{A2}$$

Proof of Proposition 1. Part (a) follows from $\phi_F > \phi_D$ and the definition in equation (6). For part (c), note that the skill composition at a firm of nationality N is

$$C_N = \frac{\ell_{Ns}}{\ell_{Ns} + \ell_{Nu}} = \frac{w_{Ns}^{\eta} \frac{\bar{L}_s}{W_s}}{w_{Ns}^{\eta} \frac{\bar{L}_s}{W_s} + w_{Nu}^{\eta} \frac{\bar{L}_u}{W_u}} = \frac{\zeta_{Ns}^{\eta}}{\zeta_{Ns}^{\eta} + \frac{\bar{L}_u / W_u}{\bar{L}_s / W_s}}, \tag{A3}$$

which only depends on N through ζ_{Ns}^{η}. Since C_N is increasing in ζ_{Ns}^{η}, then $\zeta_{Ds}^{\eta} > \zeta_{Ds}^{\eta}$ implies $C_F > C_D$, which proves part (c). Since $\zeta_{Fs} > \zeta_{Ds} \geq 1$, then $C_F > C_D$ and $C_F \log \zeta_{Fs} > C_D \log \zeta_{Ds}$, which proves part (b).

Indirect effect first-order approximations (FOA). We derive the first-order approximations around an initial equilibrium featuring a small share of employment at foreign firms. First, compute the change in foreign employment share $p = \frac{L_F}{L_F + L_D}$:

$$\Delta p = \frac{\Delta L_F}{L_F + L_D} - \frac{L_F(\Delta L_F + \Delta L_D)}{(L_F + L_D)^2} = \frac{(1-p)\Delta L_F - p\Delta L_D}{L_F + L_D} \approx \frac{\Delta L_F}{L_F + L_D} = \hat{X}.$$

It then follows that

$$\Delta \log \phi_D = \frac{\tau(\phi_F - 1)\Delta p}{1 + \tau(\phi_F - 1)p} \approx \tau(\phi_F - 1)\hat{X}. \tag{A4}$$

Indirect effect FOA for wages. From equation (5), the change in log wages at domestic firms is

$$\Delta \log w_{Dh} = \Delta \log \phi_D \approx \tau(\phi_F - 1)\hat{X}. \tag{A5}$$

Indirect effect FOA for employment. The change in log employment of skilled workers at a domestic firm $\Delta \log \ell_{Ds} = \eta \Delta \log w_{Ds} - \Delta \log W_s$ where

$$\Delta \log W_s = \frac{\ell_{Fs}}{\bar{L}_s}\Delta M_F + \eta E_s \Delta \log w_{Ds} \approx \left(\frac{C_F}{C_D}E_s + \eta\tau(\phi_F - 1)E_s\right)\hat{X}.$$

Note that we replace ΔM_F utilizing $\Delta L_F \approx \Delta M_F(\ell_{Fs} + \ell_{Fu})$.[45] Therefore,

$$\Delta \log \ell_{Ds} \approx \tau\eta(\phi_F - 1)(1 - E_s)\hat{X} - \frac{C_F}{C_D}E_s\hat{X}. \tag{A6}$$

Similarly, for the change in log employment of unskilled workers,

$$\Delta \log \ell_{Du} \approx \tau\eta(\phi_F - 1)(1 - E_u)\hat{X} - \frac{1 - C_F}{1 - C_D}E_u\hat{X}. \tag{A7}$$

The change in log total employment at a domestic firm is the mean change in log employment of both types weighted by the respective employment share. That is,

$$\Delta \log(\ell_{Ds} + \ell_{Du}) = C_D \Delta \log \ell_{Ds} + (1 - C_D)\Delta \log \ell_{Du} \approx \tau\eta(\phi_F - 1)(1 - \bar{E}_D)\hat{X} - \bar{E}_F\hat{X}, \tag{A8}$$

where $\bar{E}_N = C_N E_s + (1 - C_N)E_u$.

Indirect effect FOA for value added. From equation (1), the change in log value added at a domestic firm is $\Delta \log q_D = \Delta \log \phi_D + R_D \Delta \log \ell_{Ds} + (1 - R_D)\Delta \log \ell_{Du}$ where $R_D = \frac{\zeta_{Ds}\ell_{Ds}}{\zeta_{Ds}\ell_{Ds} + \ell_{Du}}$ is the output share of skilled workers at a domestic firm. Based on equations (A4), (A6) and (A7), we have

$$\begin{aligned}\Delta \log q_D \approx{}& \tau(\phi_F - 1)\left(1 + \eta\left[1 - R_D E_s - (1 - R_D)E_u\right]\right)\hat{X} \\ &- \left(\frac{C_F}{C_D}R_D E_s + \frac{1 - C_F}{1 - C_D}(1 - R_D)E_u\right)\hat{X}.\end{aligned} \tag{A9}$$

[45]Specifically, the first term in the equation above can be approximated as follows:

$$\frac{\ell_{Fs}}{\bar{L}_s}\Delta M_F \approx \frac{\ell_{Fs}}{\bar{L}_s}\frac{\Delta L_F}{\ell_{Fs} + \ell_{Fu}} = C_F \frac{L_F + L_D}{\bar{L}_s}\frac{\Delta L_F}{L_F + L_D} \approx \frac{C_F}{C_D}\frac{C_D L_D}{\bar{L}_s}\hat{X} = \frac{C_F}{C_D}E_s\hat{X}.$$

Indirect effect FOA for wage bill. Since $\Delta \log b_{Dh} = \Delta \log w_{Dh} + \Delta \log \ell_{Dh}$,

$$\Delta \log b_{Ds} = \left(\tau(\phi_F - 1)[1 + \eta(1 - E_s)] - \frac{C_F}{C_D} E_s \right) \hat{X},$$
$$\Delta \log b_{Du} = \left(\tau(\phi_F - 1)[1 + \eta(1 - E_u)] - \frac{1 - C_F}{1 - C_D} E_u \right) \hat{X}.$$

The change in log total wage bill at a domestic firm is the mean change in log wage bill of both types weighted by respective output share. That is,

$$\begin{aligned} \Delta \log b_D &= R_D \Delta \log b_{Ds} + (1 - R_D) \Delta \log b_{Du} \\ &\approx \tau(\phi_F - 1)\left(1 + \eta\left[1 - R_D E_s - (1 - R_D) E_u\right]\right) \hat{X} \\ &- \left(\frac{C_F}{C_D} R_D E_s + \frac{1 - C_F}{1 - C_D} (1 - R_D) E_u \right) \hat{X}. \end{aligned} \tag{A10}$$

Up to the first order, the change in log wage bill is the same as the change in log value added at a domestic firm.

Indirect effect FOA for value added per worker and wage bill per worker. From equations (A8) and (A9), the change in log value added per worker at a domestic firm is

$$\begin{aligned} \Delta \log q_D - \Delta \log(\ell_{Du} + \ell_{Ds}) &\approx \tau(\phi_F - 1)\left[1 - \eta(R_D - C_D)(E_s - E_u)\right] \hat{X} \\ &- (R_D - C_D) \left(\frac{C_F}{C_D} E_s - \frac{1 - C_F}{1 - C_D} E_u \right) \hat{X}. \end{aligned} \tag{A11}$$

The change in the wage bill per worker is identical to the right hand side of (A11).

Proof of Proposition 2. We prove Proposition 2 taking equations (7), (8), and (9) as given; they are proven above. Part (a) follows from equation (7). For part (b), consider the case in which $\tau(\phi_F - 1)$ is sufficiently large. The spillover effect from equation (8) or (9) is unbounded and increasing in ϕ_F, while the competition effect is bounded. Hence, there exists a $\bar{\phi}_F > 1$ such that $\Delta \log(\ell_{Du} + \ell_{Ds}) > 0$ and $\Delta \log q_D > 0$ for $\phi_F > \bar{\phi}_F$. Alternatively, consider the case in which E_s and E_u are sufficiently small. From equation (8) or (9), the spillover effect is decreasing in E_s and E_u, while the competition effect is increasing in E_s and E_u. When E_s and E_u are sufficiently small, the spillover effect is positive, while the competition effect approaches zero. Hence, $\Delta \log(\ell_{Du} + \ell_{Ds}) > 0$ and $\Delta \log q_D > 0$. For part (c), from equation (A11), the sign is ambiguous and depends on the magnitudes of the various terms.

Claim in footnote 16. Equation (7) uses a first-order approximation to show that the indirect effect of foreign investment has the same sign as τ. We now show that the sign of the wage effect is the same as the sign of τ without a first-order approximation:

Specifically, we show that $\frac{dw_{Dh}}{dM_F} > 0$ when $\tau > 0$ and $\frac{dw_{Dh}}{dM_F} = 0$ when $\tau = 0$. Notice that

$$\frac{dw_{Dh}}{dM_F} = \frac{\eta}{\eta+1}\frac{d\phi_D}{dM_F}\zeta_{Dh}.$$

When $\tau = 0$, $\phi_D = 1$ and $\frac{dw_{Dh}}{dM_F} = 0$. When $\tau > 0$, let

$$F(\phi_D, M_F) \equiv 1 + \tau(\phi_F - 1)\frac{L_F}{L_F + L_D} - \phi_D.$$

From the implicit function theorem, $\frac{d\phi_D}{dM_F} = -\frac{F_{M_F}}{F_{\phi_D}}$. First, we provide the elements that are used to compute F_{M_F}.

$$F_{M_F} = \tau(\phi_F - 1)\frac{\partial \frac{L_F}{L_F+L_D}}{\partial M_F} = \tau(\phi_F - 1)\frac{\frac{\partial L_F}{\partial M_F}L_D - L_F\frac{\partial L_D}{\partial M_F}}{(L_F + L_D)^2}. \tag{A12}$$

Using equations (4) and (5), we have

$$L_D = \frac{M_D(\gamma\phi_D)^\eta}{M_F(\gamma\phi_F)^\eta + M_D(\gamma\phi_D)^\eta + w_0^\eta}\bar{L}_u + \frac{M_D(\gamma\phi_D\zeta_{Ds})^\eta}{M_F(\gamma\phi_F\zeta_{Fs})^\eta + M_D(\gamma\phi_D\zeta_{Ds})^\eta + w_0^\eta}\bar{L}_s \tag{A13}$$

$$L_F = \frac{M_F(\gamma\phi_F)^\eta}{M_F(\gamma\phi_F)^\eta + M_D(\gamma\phi_D)^\eta + w_0^\eta}\bar{L}_u + \frac{M_F(\gamma\phi_F\zeta_{Fs})^\eta}{M_F(\gamma\phi_F\zeta_{Fs})^\eta + M_D(\gamma\phi_D\zeta_{Ds})^\eta + w_0^\eta}\bar{L}_s, \tag{A14}$$

where $\gamma = \frac{\eta}{\eta+1}$. Therefore,

$$\begin{aligned}
\frac{\partial L_D}{\partial M_F} &= -\frac{M_D(\gamma\phi_D)^\eta \cdot (\gamma\phi_F)^\eta}{W_u^2}\bar{L}_u - \frac{M_D(\gamma\phi_D\zeta_{Ds})^\eta \cdot (\gamma\phi_F\zeta_{Fs})^\eta}{W_s^2}\bar{L}_s, \\
\frac{\partial L_F}{\partial M_F} &= \frac{[M_D(\gamma\phi_D)^\eta + w_0^\eta] \cdot (\gamma\phi_F)^\eta}{W_u^2}\bar{L}_u + \frac{[M_D(\gamma\phi_D\zeta_{Ds})^\eta + w_0^\eta] \cdot (\gamma\phi_F\zeta_{Fs})^\eta}{W_s^2}\bar{L}_s.
\end{aligned}$$

We see that $\frac{\partial L_D}{\partial M_F} < 0$ and $\frac{\partial L_F}{\partial M_F} > 0$. This implies that $F_{M_F} > 0$.

Next, we provide the elements that are used to compute F_{ϕ_D}:

$$F_{\phi_D} = \tau(\phi_F - 1)\frac{\partial \frac{L_F}{L_F+L_D}}{\partial \phi_D} - 1 = \tau(\phi_F - 1)\frac{\frac{\partial L_F}{\partial \phi_D}L_D - L_F\frac{\partial L_D}{\partial \phi_D}}{(L_F + L_D)^2} - 1, \tag{A15}$$

where

$$\frac{\partial L_D}{\partial \phi_D} = \frac{\gamma\eta M_D(\gamma\phi_D)^{\eta-1}[M_F(\gamma\phi_F)^\eta + w_0^\eta]}{W_u^2}\bar{L}_u + \frac{\gamma\eta\zeta_{Ds}M_D(\gamma\phi_D\zeta_{Ds})^{\eta-1}[M_F(\gamma\phi_F\zeta_{Fs})^\eta + w_0^\eta]}{W_s^2}\bar{L}_s$$
$$\frac{\partial L_F}{\partial \phi_D} = -\frac{\gamma\eta M_D(\gamma\phi_D)^{\eta-1}M_F(\gamma\phi_F)^\eta}{W_u^2}\bar{L}_u - \frac{\gamma\eta\zeta_{Ds}M_D(\gamma\phi_D\zeta_{Ds})^{\eta-1}M_F(\gamma\phi_F\zeta_{Fs})^\eta}{W_s^2}\bar{L}_s.$$

We see that $\frac{\partial L_D}{\partial \phi_D} > 0$ and $\frac{\partial L_F}{\partial \phi_D} < 0$. This implies that $F_{\phi_D} < 0$ and $\frac{d\phi_D}{dM_F} = -\frac{F_{M_F}}{F_{\phi_D}} > 0$. Therefore, we have $\frac{dw_{Dh}}{dM_F} = \frac{\eta}{\eta+1}\frac{d\phi_D}{dM_F}\zeta_{Dh} > 0$ when $\tau > 0$.

A.4 Model with Many Skill and Firm Types

We next provide a model with an arbitrary number firm and worker types as well as many foreign nationalities.

Environment. We assume there is a large set of locations in the U.S. All regions are trading frictionless within the U.S. and workers are immobile across locations. We focus on the outcomes in one particular location and, to simplify notation, omit the locations subscript. Let $N \in \{D, 1, ..., \bar{N}\}$ denote the firm country of origin, where $N = D$ is domestic and $N \geq 1$ indexes the foreign nationalities. Let $N(j)$ denote the nationality of firm j. Denote by M_N the number of firms of nationality N. Let $h(i)$ denote the skill level of a worker i. Denote by L_{Nh} the number of employees with skill level h, and $L_N = \sum_h L_{Nh}$ is the total number of employees at firms of nationality N. Each region is equipped with $\bar{L}_h$ potential employees of skill type h.

Preferences and labor supply. These are unchanged from the main text, except there are more values of h; see equation (4). We normalize the minimum value of h to 1 without loss of generality.

Technology. Each firm produces a homogeneous good q that is freely traded, where the price is normalized to 1. A firm produces using technology,

$$q_j(\{\ell_h\}_h) = \phi_j \sum_h h^{\theta_j} \ell_h, \tag{A16}$$

where ϕ_j is firm-specific TFP and θ_j is the firm-specific skilled-labor-augmenting productivity parameter. If $\theta_j > 1$, the firm-specific productivity of labor is increasing at an increasing rate in skill level h; if $0 < \theta_j < 1$, the firm-specific productivity of labor is increasing at a decreasing rate in skill level h.

Equilibrium wages. Given the production function in (A16) and labor supply in equation (4), equilibrium wages are given by

$$w_{ij} = \frac{\eta}{\eta + 1} \phi_j h_i^{\theta_j} \tag{A17}$$

Defining $\mu \equiv \log \frac{\eta}{\eta+1}$, $\psi_j \equiv \log \phi_j$, and $x_i \equiv \log h_i$, the equilibrium log wage is

$$\log w_{ij} = \mu + \psi_j + \theta_j x_i. \tag{A18}$$

From this expression, the mean difference in log wages for firms with foreign nationality N relative to domestic firms D is

$$\underbrace{\mathbb{E}[\log w_{ij}|N(j)=N]-\mathbb{E}[\log w_{ij}|N(j)=D]}_{\text{Total country } N \text{ wage differential}}=\underbrace{\mathbb{E}[x_i|N(j)=N]-\mathbb{E}[x_i|N(j)=D]}_{\text{Skill composition difference for country } N}$$
$$+\underbrace{\mathbb{E}[\psi_j|N(j)=N]-\mathbb{E}[\psi_j|N(j)=D]}_{\text{Low-skill country } N \text{ premium}}+\underbrace{\mathbb{E}[(\theta_j-1)x_i|N(j)=N]-\mathbb{E}[(\theta_j-1)x_i|N(j)=D]}_{\text{Additional country } N \text{ premium due to skill augmentation}} \tag{A19}$$

Suppose that there are $\bar{F}$ ($\bar{D}$) different types of foreign (domestic) firms. Denote the set of foreign (domestic) firm types by $\mathcal{F}$ ($\mathcal{D}$). Foreign firms with type $f \in \mathcal{F}$ are characterized by the pair (ϕ_f, θ_f). The type of domestic firms is characterized by a pair of base productivities, $(\tilde{\phi}_d, \tilde{\theta}_d)$, with $d \in \mathcal{D}$. The *ex post* productivities at a type-d domestic firm are determined as follows:

$$\phi_d = \tilde{\phi}_d + \tau \sum_f \frac{L_f}{L_D + L_F}(\phi_f - \tilde{\phi}_d) \tag{A20}$$

$$\theta_d = \tilde{\theta}_d + \tau \sum_f \frac{L_f}{L_D + L_F}(\theta_f - \tilde{\theta}_d), \tag{A21}$$

where $L_F = \sum_f \sum_h L_{fh}$ is the total employment at foreign firms.

In addition, denote the mass of type-f foreign firms with nationality N as M_f^N and the total mass of firms with nationality N as $M^N = \sum_f M_f^N$. Without loss of generality, we order (ϕ_f, θ_f) such that θ_f is increasing in f.

Assumption 1 *For any two countries N and N', either $\frac{\sum_{f>\bar{f}} M_f^N}{M^N} \geq \frac{\sum_{f>\bar{f}} M_f^{N'}}{M^{N'}}$ or $\frac{\sum_{f>\bar{f}} M_f^N}{M^N} \leq \frac{\sum_{f>\bar{f}} M_f^{N'}}{M^{N'}}$ holds for all $\bar{f}$. In addition, $\frac{\sum_{f>\bar{f}} M_f^N}{M^N} \geq \frac{\sum_{d>\bar{f}} M_d^D}{M^D}$ holds for all N and $\bar{f}$.*

Based on Assumption 1, we are able to rank foreign countries by their respective average skilled-labor-augmenting productivity, $\bar{\theta}^N = \sum_f \frac{M_f^N}{M^N}\theta_f$. Note that given two foreign countries N and N', $\frac{\sum_{f>\bar{f}} M_f^N}{M^N} \geq \frac{\sum_{f>\bar{f}} M_f^{N'}}{M^{N'}}$ for all $\bar{f}$ if and only if $\bar{\theta}^N \geq \bar{\theta}^{N'}$ for all increasing sequences of θ_f.

Lemma 1 *$\sum_{h>\bar{h}} C_{fh}$ is non-decreasing in θ_f for all $\bar{h}$.*

Proof. Define the share of workers with skill level h in a type-f foreign firm as $C_{fh} = \frac{\ell_{fh}}{\sum_g \ell_{fg}}$. When $\bar{h} \geq \bar{H}$, $\sum_{h>\bar{H}} C_{fh} = 0$ for all θ_f. When $0 < \bar{h} < \bar{H}$,

$$\sum_{h>\bar{h}} C_{fh} = \frac{1}{1 + \frac{\sum_{g\leq\bar{h}} g^{\eta\theta_f}\bar{L}_g/W_g}{\sum_{h>\bar{h}} h^{\eta\theta_f}\bar{L}_h/W_h}}. \tag{A22}$$

Let $G(\theta_f, \bar{h}) = \frac{\sum_{g\leq \bar{h}} g^{\eta\theta_f} \bar{L}_g/W_g}{\sum_{h>\bar{h}} h^{\eta\theta_f} \bar{L}_h/W_h}$, then

$$\begin{aligned}\frac{\partial G(\theta_f, \bar{h})}{\partial \theta_f} &= \frac{den \cdot \eta \sum_{g\leq \bar{h}} g^{\eta\theta_f} \log g \cdot \bar{L}_g/W_g - num \cdot \eta \sum_{h>\bar{h}} h^{\eta\theta_f} \log h \cdot \bar{L}_h/W_h}{den^2} \\ &< \eta \log \bar{h} \frac{num \cdot den - num \cdot den}{den^2} = 0.\end{aligned}$$

In this case, $\sum_{h>\bar{h}} C_{fh}$ is strictly increasing in θ_f. ■

Proposition 3 (Direct effects with many foreign countries and skill types) *Suppose that Assumption 1 holds, then in equation (A19),*

(a) "Skill composition difference for country N" and "Additional country N premium due to skill augmentation" are positive.

(b) "Skill composition difference for country N" and "Additional country N premium due to skill augmentation" are increasing in $\bar{\theta}_N$.

Proof. For the function $H = \log h$ or $H = (\theta_f - 1)\log h$,

$$\mathbb{E}[H|N(j) = N] - \mathbb{E}[H|N(j) = N'] = \sum_{f\in\mathcal{F}} \frac{M_f^N}{M^N} \sum_{h\in\mathcal{H}} C_{fh} H - \sum_{f\in\mathcal{F}} \frac{M_f^{N'}}{M^{N'}} \sum_{h\in\mathcal{H}} C_{fh} H$$

Suppose that $\bar{\theta}_N \geq \bar{\theta}_{N'}$, which implies that $\mathbb{E}[H|N(j) = N] - \mathbb{E}[H|N(j) = N'] \geq 0$ from Assumption 1. In particular, let $N' = D$, and part (a) can be proved based on Assumption 1 and Lemma 1. ■

Two-way fixed effect model. Consider a special case in which $\theta_j \equiv \bar{\theta}$, so that the skill-augmenting technology is homogeneous. Furthermore, we assume x_i can be decomposed as $x_i = \tilde{x}_i + \chi_i'\tilde{\beta}$, where $\tilde{x}_i$ is unobserved to the econometrician and $\chi_i'\tilde{\beta}$ is observed to the econometrician. Then, we can write $\theta_j x_i = \bar{\theta}\left(\tilde{x}_i + \chi_i'\tilde{\beta}\right) = \grave{x}_i + \chi_i'\beta$ where $\grave{x}_i \equiv \bar{\theta}\tilde{x}_i$ and $\beta \equiv \bar{\theta}\tilde{\beta}$. This implies that $\log w_{ij} = \phi_j + \grave{x}_i + \chi_i'\beta$.

Indirect effect first-order approximations (FOA). First, we derive the FOA of $p_f = \frac{L_f}{L_D + L_F}$, the share of workers employed in a type-f foreign firm. Throughout this section, we conduct the FOA around an initial equilibrium in which the employment share at foreign-owned firms is small:

$$\Delta p_f = \frac{\Delta L_f - \frac{L_f}{L_F}\Delta L_F}{L_D + L_F} + \frac{L_f}{L_F}\Delta p \approx \frac{L_f}{L_F}\hat{X}.$$

Indirect effect FOA for wages.

$$\Delta \log w_{dh} = \Delta \log \phi_d + \Delta\theta_d \log h,$$

where

$$\Delta \log \phi_d = \frac{\tau \sum_{f\in\mathcal{F}} \Delta p_f(\phi_f - \tilde{\phi}_d)}{\tilde{\phi}_d + \tau \sum_{f\in\mathcal{F}} p_f(\phi_f - \tilde{\phi}_d)} \approx \tau \sum_{f\in\mathcal{F}} \frac{\phi_f - \tilde{\phi}_d}{\tilde{\phi}_d} \frac{L_f}{L_F} \hat{X}$$
$$\Delta\theta_d = \tau \sum_{f\in\mathcal{F}} (\theta_f - \tilde{\theta}_d)\Delta p_f \approx \tau \sum_{f\in\mathcal{F}} (\theta_f - \tilde{\theta}_d) \frac{L_f}{L_F} \hat{X}.$$

Therefore,

$$\Delta \log w_{dh} \approx \tau \sum_{f\in\mathcal{F}} \left[\frac{\phi_f - \tilde{\phi}_d}{\tilde{\phi}_d} + (\theta_f - \tilde{\theta}_d) \log h \right] \frac{L_f}{L_F} \hat{X},$$

and the expected change in wage across all domestic firm types and worker skill levels is

$$\mathbb{E}[\Delta \log w_{dh}] = \tau \sum_{d\in\mathcal{D}} \frac{M_d^D}{M^D} \sum_{h\in\mathcal{H}} C_{dh} \sum_{f\in\mathcal{F}} \left[\frac{\phi_f - \tilde{\phi}_d}{\tilde{\phi}_d} + (\theta_f - \tilde{\theta}_d) \log h \right] \frac{L_f}{L_F} \hat{X}. \tag{A23}$$

Indirect effect FOA for employment. Since

$$\begin{aligned} \Delta \log W_h &= \sum_{f\in\mathcal{F}} \frac{\ell_{fh}}{\bar{L}_h} \Delta M_f + \eta \sum_{d\in\mathcal{D}} E_{dh} \Delta \log w_{dh} \\ &\approx \sum_{f\in\mathcal{F}} C_{fh} \frac{L_D}{\bar{L}_h} \frac{L_f}{L_F} \hat{X} + \tau\eta \sum_{d\in\mathcal{D}} E_{dh} \sum_{f\in\mathcal{F}} \left[\frac{\phi_f - \tilde{\phi}_d}{\tilde{\phi}_d} + (\theta_f - \tilde{\theta}_d) \log h \right] \frac{L_f}{L_F} \hat{X}, \end{aligned}$$

then

$$\begin{aligned} \Delta \log \ell_{dh} &= \eta \Delta \log w_{dh} - \Delta \log W_h \\ &\approx \eta \left[(1 - E_{dh}) \Delta \log w_{dh} - \sum_{g\in\mathcal{D}\backslash d} E_{gh} \Delta \log w_{gh} \right] - \sum_{f\in\mathcal{F}} C_{fh} \frac{L_D}{\bar{L}_h} \frac{L_f}{L_F} \hat{X}. \end{aligned}$$

Therefore,

$$
\begin{aligned}
&\Delta \log \left(\sum_{h \in \mathcal{H}} \ell_{dh} \right) = \sum_{h \in \mathcal{H}} C_{dh} \Delta \log \ell_{dh} \\
\approx &\eta \sum_{h \in \mathcal{H}} C_{dh} \left[(1 - E_{dh}) \Delta \log w_{dh} - \sum_{g \in \mathcal{D} \setminus d} E_{gh} \Delta \log w_{gh} \right] - \frac{L_D}{L_d} \sum_{h \in \mathcal{H}} \sum_{f \in \mathcal{F}} C_{fh} E_{dh} \frac{L_f}{L_F} \hat{X},
\end{aligned}
$$

and the expected change in employment across all domestic firm types is

$$
\begin{aligned}
&\mathbb{E} \left[\Delta \log \left(\sum_{h \in \mathcal{H}} \ell_{dh} \right) \right] = \sum_{d \in \mathcal{D}} \frac{M_d^D}{M^D} \sum_{h \in \mathcal{H}} C_{dh} \Delta \log \ell_{dh} \\
\approx &\eta \sum_{d \in \mathcal{D}} \frac{M_d^D}{M^D} \sum_{h \in \mathcal{H}} C_{dh} \left[(1 - E_{dh}) \Delta \log w_{dh} - \sum_{g \in \mathcal{D} \setminus d} E_{gh} \Delta \log w_{gh} \right] - \sum_{d \in \mathcal{D}} \frac{M_d^D}{M^D} \frac{L_D}{L_d} \sum_{h \in \mathcal{H}} \sum_{f \in \mathcal{F}} C_{fh} E_{dh} \frac{L_f}{L_F} \hat{X}.
\end{aligned}
\tag{A24}
$$

Indirect effect FOA for wage bill. Since

$$
\Delta \log b_{dh} = \Delta \log w_{dh} + \Delta \log \ell_{dh} \approx [1 + \eta (1 - E_{dh})] \Delta \log w_{dh} - \eta \sum_{g \in \mathcal{D} \setminus d} E_{gh} \Delta \log w_{gh} - \sum_{f \in \mathcal{F}} C_{fh} \frac{L_D}{\bar{L}_h} \frac{L_f}{L_F} \hat{X},
$$

then

$$
\begin{aligned}
&\Delta \log \left(\sum_{h \in \mathcal{H}} b_{dh} \right) = \sum_{h \in \mathcal{H}} R_{dh} \Delta \log b_{dh} \\
\approx &\sum_{h \in \mathcal{H}} R_{dh} \left([1 + \eta (1 - E_{dh})] \Delta \log w_{dh} - \eta \sum_{g \in \mathcal{D} \setminus d} E_{gh} \Delta \log w_{gh} \right) - \sum_{h \in \mathcal{H}} R_{dh} \frac{L_D}{L_{dh}} \sum_{f \in \mathcal{F}} C_{fh} E_{dh} \frac{L_f}{L_F} \hat{X},
\end{aligned}
$$

and the expected change in wage bill across all domestic firm types is

$$
\begin{aligned}
&\mathbb{E} \left[\Delta \log \left(\sum_{h \in \mathcal{H}} b_{dh} \right) \right] = \sum_{d \in \mathcal{D}} \frac{M_d^D}{M^D} \sum_{h \in \mathcal{H}} R_{dh} \Delta \log b_{dh} \\
\approx &\sum_{d \in \mathcal{D}} \frac{M_d^D}{M^D} \sum_{h \in \mathcal{H}} R_{dh} \left([1 + \eta (1 - E_{dh})] \Delta \log w_{dh} - \eta \sum_{g \in \mathcal{D} \setminus d} E_{gh} \Delta \log w_{gh} \right) \\
&- \sum_{d \in \mathcal{D}} \frac{M_d^D}{M^D} \sum_{h \in \mathcal{H}} R_{dh} \frac{L_D}{L_{dh}} \sum_{f \in \mathcal{F}} C_{fh} E_{dh} \frac{L_f}{L_F} \hat{X}.
\end{aligned}
\tag{A25}
$$

Indirect effect FOA for value added.

$$
\begin{aligned}
\Delta \log q_d =& \Delta \log \phi_d + \Delta \log \left(\sum_{h \in \mathcal{H}} h^{\theta_d} \ell_{dh} \right) = \Delta \log \phi_d + \sum_{h \in \mathcal{H}} R_{dh} (\Delta \theta_d \log h + \Delta \log \ell_{dh}) \\
\approx & \tau \sum_{f \in \mathcal{F}} \frac{\phi_f - \tilde{\phi}_d}{\tilde{\phi}_d} \frac{L_f}{L_F} \hat{X} + \sum_{h \in \mathcal{H}} R_{dh} \left(\tau \log h \sum_{f \in \mathcal{F}} (\theta_f - \tilde{\theta}_d) \frac{L_f}{L_F} \hat{X} + \Delta \log \ell_{dh} \right) \\
\approx & \sum_{h \in \mathcal{H}} R_{dh} (\Delta \log w_{dh} + \Delta \log \ell_{dh}) = \sum_{h \in \mathcal{H}} R_{dh} \Delta \log b_{dh} = \Delta \log \left(\sum_{h \in \mathcal{H}} b_{dh} \right)
\end{aligned}
$$

Therefore,

$$
\mathbb{E} \left[\Delta \log q_d \right] = \mathbb{E} \left[\Delta \log \left(\sum_{h \in \mathcal{H}} b_{dh} \right) \right]
$$

Indirect effect FOA for value added per worker.

$$
\begin{aligned}
& \Delta \log q_d - \Delta \log \left(\sum_{h \in \mathcal{H}} \ell_{dh} \right) \\
\approx & \sum_{h \in \mathcal{H}} R_{dh} \left([1 + \eta (1 - E_{dh})] \Delta \log w_{dh} - \eta \sum_{g \in \mathcal{D} \setminus d} E_{gh} \Delta \log w_{gh} \right) - \sum_{h \in \mathcal{H}} R_{dh} \frac{L_D}{L_{dh}} \sum_{f \in \mathcal{F}} C_{fh} E_{dh} \frac{L_f}{L_F} \hat{X} \\
& - \eta \sum_{h \in \mathcal{H}} C_{dh} \left[(1 - E_{dh}) \Delta \log w_{dh} - \sum_{g \in \mathcal{D} \setminus d} E_{gh} \Delta \log w_{gh} \right] - \frac{L_D}{L_d} \sum_{h \in \mathcal{H}} \sum_{f \in \mathcal{F}} C_{fh} E_{dh} \frac{L_f}{L_F} \hat{X} \\
= & \sum_{h \in \mathcal{H}} \left([R_{dh} + \eta (R_{dh} - C_{dh})(1 - E_{dh})] \Delta \log w_{dh} - \eta (R_{dh} - C_{dh}) \sum_{g \in \mathcal{D} \setminus d} E_{gh} \Delta \log w_{gh} \right) \\
& - \sum_{h \in \mathcal{H}} (R_{dh} - C_{dh}) \frac{L_D}{L_{dh}} \sum_{f} C_{fh} E_{dh} \frac{L_f}{L_F} \hat{X} \qquad \text{(A26)}
\end{aligned}
$$

Proposition 4 (Indirect effects with many foreign countries and skill types) *If $\min_{f \in \mathcal{F}} \phi_f \geq \max_{d \in \mathcal{D}} \tilde{\phi}_d$, $\min_{f \in \mathcal{F}} \theta_f \geq \max_{d \in \mathcal{D}} \tilde{\theta}_d$, and foreign firms have positive spillovers onto domestic firms (i.e., $\tau > 0$), then — up to a first-order approximation around an initial equilibrium featuring a small share of employment at foreign firms — an increase in the share of employment at foreign firms causes*

(a) A positive effect on mean wages at domestic firms;

(b) A positive effect on mean employment, mean wage bill, and mean value added at domestic firms if E_{dh} is sufficiently small for all $d \in \mathcal{D}$ and $h \in \mathcal{H}$;

(c) An ambiguous effect on mean value added per worker at domestic firms.

Proof. Part (a) follows from equation (A23). For part (b), when E_{dh} is sufficiently small for all d and h, the spillover effect is decreasing in E_{dh}, while the competition effect is increasing in E_{dh} in equations (A24) and (A25). When E_{dh} is sufficiently small for all d and h, the spillover effect is positive, while the competition effect approaches zero. Hence, $\mathbb{E}[\Delta \log(\sum_{h\in\mathcal{H}} \ell_{dh})] > 0$ and $\mathbb{E}[\Delta \log(\sum_{h\in\mathcal{H}} b_{dh})] = \mathbb{E}[\Delta \log q_D] > 0$. For part (c), from equation (A26), the sign is ambiguous and depends on the magnitudes of the various terms. ■

A.5 Direct Effects: Evidence from Movers

As an alternate to equation (10), we use a difference-in-differences design for workers that move across firms. Here, we allow for asymmetric wage changes between workers that move from domestic to foreign firms and those that move the other way. However, as in the theory, domestic and foreign are the only firm types. By looking at within-worker differences in wages, we remove the worker-specific time-invariant wage level.

To implement the difference-in-differences design for movers across firms, we define the following indicator variables:

$M_{i,t,DF}$: worker i moving from a domestic firm in $t-1$ to a foreign firm in t;

$M_{i,t,FD}$: worker i moving from a foreign firm in $t-1$ to a domestic firm in t;

$M_{i,t,DD}$: worker i moving from a domestic firm in $t-1$ to a domestic firm in t; and

$M_{i,t,FF}$: worker i moving from a foreign firm in $t-1$ to a foreign firm in t.

Equipped with these indicator variables summarizing the workers job transition status, we estimate the following regression model:

$$\log w_{i,t+1} - \log w_{i,t-2} = \beta_{FF} M_{i,t,FF} + \beta_{FD} M_{i,t,FD} + \beta_{DF} M_{i,t,DF} + \mu_{cz(i),t+1} + \nu_{ind(i),t+1} + \tilde{\mu}_{cz(i),t-2} + \tilde{\nu}_{ind(i),t-2} + \epsilon_{i,t}, \quad \text{(A27)}$$

where we omit $M_{i,t,DD}$ so that domestic to domestic moves serve as the control group. The regression controls consist of the industry-year fixed effects (both for the industry in year $t+1$ and in year $t-2$), commuting-zone-year fixed effects (both for the commuting zone in year $t+1$ and in year $t-2$), and a polynomial in age (to remove age-related wage growth). The sample consists only of workers that are in different firms in $t+1$ and $t-2$. We do not measure the outcome during the intermediate years $t-1$ and t because earnings may account for partial years of employment only while the worker is in the process of moving.

The main results are presented in Appendix Table A2. In the baseline specification, we find that moving from a domestic to a foreign firm is associated with a 6 percent increase in wages (relative to wage growth for workers who move between domestic firms), while a 4 percent decrease in wages is associated with moving from a foreign to a domestic firm (either could be interpreted as an estimate of the average foreign firm premium).[46] Appendix Figure A.4 provides suggestive visual evidence that the effects are not driven by trends that existed prior to the moves. The slight asymmetry in effects is consistent with firm-worker interactions, as in Subsection A.6.

We consider three sample restrictions. First, we restrict the sample of domestic firms to only include non-multinationals. We find that the estimates become stronger at 8 percent

[46] Similar results for job movers are found by Martins and Esteves (2015) in Brazil.

when moving to a foreign firm and 6 percent when moving from a foreign firm. Second, we further restrict the sample to workers that separate in a mass layoff event. To do so, we restrict the sample to firms that had at least 10 workers in the first two years and 30 percent of those workers move to a different firm in the latter two years.[47] We find a 6 percent wage gain when moving from a domestic to a foreign firm and a 5 percent wage loss when moving in the reverse direction. Third, we restrict the domestic firms to only include multinationals. We find a 0 percent wage gain when moving from a domestic multinational to a foreign firm and a 1 percent wage gain when moving in the reverse direction. This is consistent with our finding above that there is little to no difference in the average premiums of domestic and foreign multinationals.

Table A2: Difference-in-Differences Estimates of the Average Foreign Firm Premium

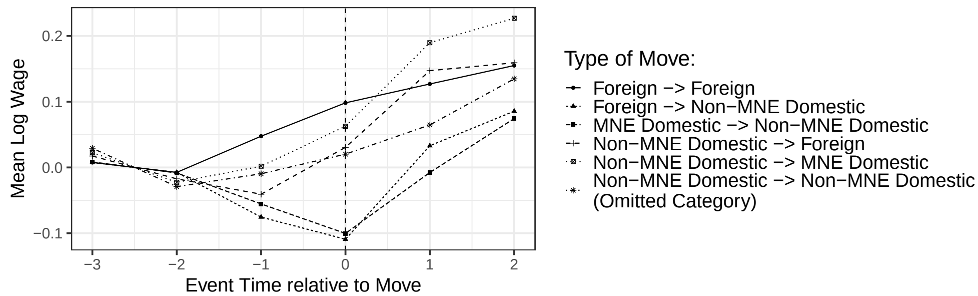
	(1)	(2)	(3)
Type of Move:			
Domestic to Foreign	0.078***	0.059***	-0.002
	(0.002)	(0.003)	(0.004)
	(N=242,207)	(N=126,178)	(N=48,795)
Foreign to Domestic	-0.056***	-0.052***	0.014***
	(0.002)	(0.004)	(0.004)
	(N=172,896)	(N=46,729)	(N=37,966)
Foreign to Foreign	0.020***	0.042***	0.006*
	(0.003)	(0.004)	(0.003)
	(N=246,192)	(N=128,396)	(N=246,192)
Domestic to Domestic	0	0	0
(Omitted Category)	(N=7,900,458)	(N=3,290,933)	(N=223,424)
Specification Details:			
Domestic Firms Restriction	Exclude MNE	Exclude MNE	Only include MNE
Type of Separation	All	Mass Layoff	All

Notes: This table presents the main effects of interest in the saturated difference-in-differences specification described in the text. The sample consists of only workers who were employed for two straight years at one firm followed by two straight years at a different firm. In column (1), we restrict the sample to domestic non-multinationals and foreign firms. In column (2), we restrict the sample to domestic non-multinationals and foreign firms and also restrict the sample to workers who separated from a firm as part of a mass layoff. In column (3), we restrict the sample to domestic multinationals and foreign firms. Standard errors are clustered by commuting-zone-year.

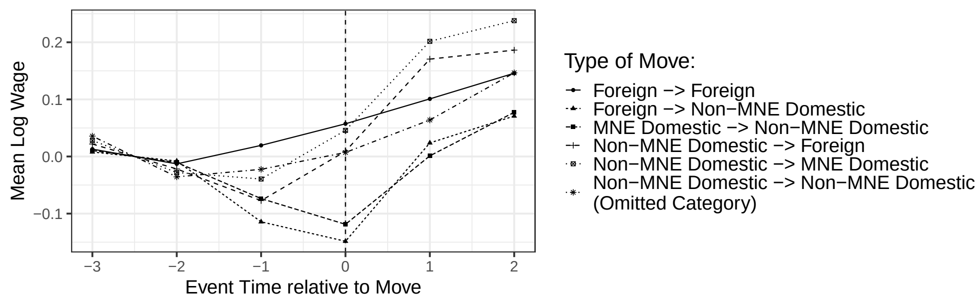
[47] We follow Yagan (2019) in using a 30 percent separation rate threshold when defining mass layoffs.

Figure A.4: Event Study for Movers to and from Foreign Firms

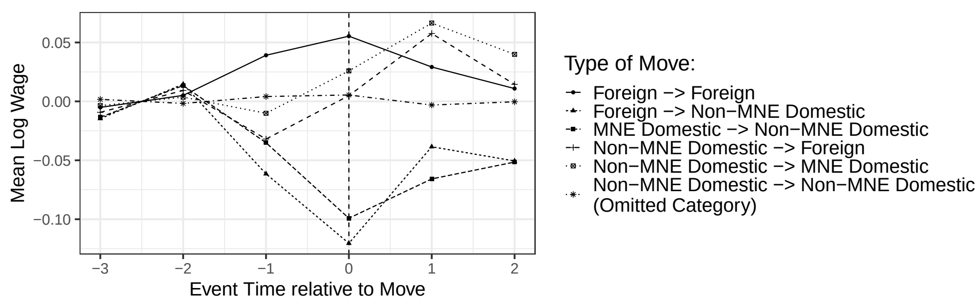
(a) Full Sample: Raw Log Wage (not controlling for age, industry-year, or CZ-year)



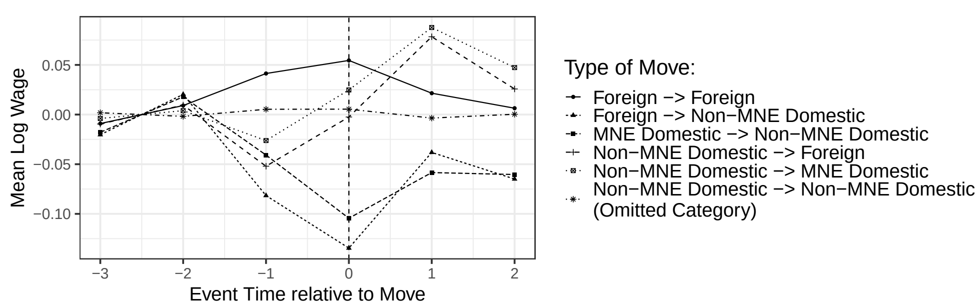
(b) Mass Layoff Sample: Raw Log Wage (not controlling for age, industry-year, or CZ-year)



(c) Full Sample: Residual Log Wage (controlling for age, industry-year, or CZ-year)



(d) Mass Layoff Sample: Residual Log Wage (controlling for age, industry-year, or CZ-year)



Notes: This figure plots mean log wages for the sample of workers that move firms. Mean log wages are normalized to be zero on average over event times -3 and -2. This figure considers two samples: Full Sample (subfigures a and c), which indicates all workers satisfying the employment spell requirements, and Mass Layoff Sample (subfigures b and d), indicating workers at firms that lost 30 percent or more of their employees in a given year. It provides two measures of the mean log wage: Raw Log Wage (subfigures a and b), indicating the unadjusted log wage, and Residual Log Wage (subfigures c and d), indicating the log wage residuals from a regression on an age polynomial, commuting-zone-year fixed effects, and industry-year fixed effects.

A.6 Direct Effects: Extension with Firm-Worker Interactions

Identification of θ. The model of firm-specific skill-augmenting productivity presented in Appendix A.4 implies the equilibrium wage-setting regression equation,

$$\log w_{i,t} = \psi_{j(i,t)} + \theta_j x_i + \epsilon_{i,t}.$$

Firms are grouped into k types, where all firms of the same type have the same ψ and θ. To understand how this model is identified, consider workers moving from firm type A at time t to firm type B at time $t+1$. Denoting this set of movers by $A_t \to B_{t+1}$, we consider the identifying content of the estimator $\widehat{\theta}_{A,B}$ defined by

$$\widehat{\theta}_{A,B} \equiv \frac{\mathbb{E}[\log w_{i,t+1}|A_t \to B_{t+1}] - \mathbb{E}[\log w_{i,t}|B_t \to A_{t+1}]}{\mathbb{E}[\log w_{i,t+1}|B_t \to A_{t+1}] - \mathbb{E}[\log w_{i,t}|A_t \to B_{t+1}]}.$$

The identification argument follows Bonhomme et al. (2019). First, notice that

$$\mathbb{E}[\log w_{i,t}|A_t \to B_{t+1}] = \psi_A + \mathbb{E}[\theta_A x_i + \epsilon_{i,t}|A_t \to B_{t+1}]$$
$$\mathbb{E}[\log w_{i,t+1}|A_t \to B_{t+1}] = \psi_B + \mathbb{E}[\theta_B x_i + \epsilon_{i,t+1}|A_t \to B_{t+1}].$$

Second, we see that $\widehat{\theta}_{A,B}$ does not involve ψ, as it simplifies to

$$\begin{aligned}\widehat{\theta}_{A,B} &= \frac{(\psi_B + \mathbb{E}[\theta_B x_i + \epsilon_{i,t+1}|A_t \to B_{t+1}]) - (\psi_B + \mathbb{E}[\theta_B x_i + \epsilon_{i,t}|B_t \to A_{t+1}])}{(\psi_A + \mathbb{E}[\theta_A x_i + \epsilon_{i,t+1}|B_t \to A_{t+1}]) - (\psi_A + \mathbb{E}[\theta_A x_i + \epsilon_{i,t}|A_t \to B_{t+1}])} \\ &= \frac{\theta_B\left(\mathbb{E}[x_i|A_t \to B_{t+1}] - \mathbb{E}[x_i|B_t \to A_{t+1}]\right) + \left(\mathbb{E}[\epsilon_{i,t+1}|A_t \to B_{t+1}] - \mathbb{E}[\epsilon_{i,t}|B_t \to A_{t+1}]\right)}{\theta_A\left(\mathbb{E}[x_i|B_t \to A_{t+1}] - \mathbb{E}[x_i|A_t \to B_{t+1}]\right) + \left(\mathbb{E}[\epsilon_{i,t+1}|B_t \to A_{t+1}] - \mathbb{E}[\epsilon_{i,t}|A_t \to B_{t+1}]\right)}.\end{aligned}$$

Third, under the assumption that workers endogenously move across firms based only on (x, ψ, θ) but do not select moves based on the measurement error ϵ, it follows that the expectation of ϵ is zero conditional on $A_t \to B_{t+1}$ or $B_t \to A_{t+1}$. Therefore,

$$\widehat{\theta}_{A,B} = \frac{\theta_B\left(\mathbb{E}[x_i|A_t \to B_{t+1}] - \mathbb{E}[x_i|B_t \to A_{t+1}]\right) + 0}{\theta_A\left(\mathbb{E}[x_i|B_t \to A_{t+1}] - \mathbb{E}[x_i|A_t \to B_{t+1}]\right) + 0} = \frac{-\theta_B}{\theta_A}, \tag{A28}$$

where the second equality requires $\mathbb{E}[x_i|B_t \to A_{t+1}] \neq \mathbb{E}[x_i|A_t \to B_{t+1}]$. This means that different firm types must attract different skill types, which is consistent with our model and empirical findings.

Thus, for any two firm types A and B, the estimator $\widehat{\theta}_{A,B}$ identifies θ_B/θ_A. Normalizing the first firm type to $\theta = 1$, which is without loss of generality since we are only interested in relative differences, this estimator identifies θ_j for each firm j.

Estimation of θ and ψ. While the derivation above helps to understand how θ_j is identified separately from ψ_j, in practice we simultaneously estimate (ψ_j, θ_j) using the following moment equation:

$$\mathbb{E}\left[\left(\frac{\log w_{i,t+1}}{\theta_{j'}} - \frac{\psi_{j'}}{\theta_{j'}}\right) - \left(\frac{\log w_{i,t}}{\theta_j} - \frac{\psi_j}{\theta_j}\right) \big| j(i,t) = j, j(i,t+1) = j'\right] = 0. \tag{A29}$$

With $k = 10$ firm types, there are 90 such moment equations with $j \neq j'$ that we can use to estimate the 20 parameters, so this is an over-identified system of equations for (ψ_j, θ_j). In practice, we use the limited information maximum likelihood (LIML) estimator of Bonhomme et al. (2019) and the R software implementation provided by their paper.

Identification and estimation of x. The final step is to identify x_i. To do so, we rearrange the wage equation and take the expectation across time periods for a given worker i:

$$x_i = \mathbb{E}\left[\frac{\log w_{i,t} - \psi_{j(i,t)}}{\theta_{j(i,t)}}\Big| i\right] - \mathbb{E}\left[\frac{\epsilon_{i,t}}{\theta_{j(i,t)}}\Big| i\right]. \tag{A30}$$

Again using that $j(i,t)$ is chosen exogenously of the measurement error $\epsilon_{i,t}$, the second expectation term is zero, so x_i is identified if (ψ_j, θ_j) are identified. In practice, we estimate x_i by simply replacing this moment condition with its sample counterpart

$$x_i = \frac{1}{|\mathcal{T}_i|} \sum_{t \in \mathcal{T}_i} \frac{\log w_{i,t} - \psi_{i,j(i,t)}}{\theta_{i,j(i,t)}}, \tag{A31}$$

where $\mathcal{T}_i$ denotes the set of time periods during which individual i is employed, and the right-hand side uses the estimates of (ψ_j, θ_j) discussed above. See also Lamadon et al. (2020) for related discussion.

Clustering firms into types. We demonstrated above that, given the k firm types, we identify (x_i, ψ_j, θ_j). To determine the assignment of firms to types, we follow Bonhomme et al. (2019) in grouping firms into k clusters using the k-means algorithm applied to the within-firm distribution of log wages. Let $c(j)$ denote the cluster of firm j, where $c = 1, 2, ..., k$. To determine the clusters, we solve the weighted k-means problem

$$\min_{c(1),\ldots,c(J),H_1,\ldots,H_k} \sum_{j=1}^{J} N_j \int \left(\widehat{F}_j(w) - H_{c(j)}(w)\right)^2 d\mu(w),$$

where $\widehat{F}_j(w)$ denotes the empirical cumulative distribution function (CDF) of log wages within firm j, N_j is the total number of workers in firm j, μ is the measure corresponding to a grid of quantiles at which the CDF is evaluated, and H_c is a candidate CDF of the log wages in cluster c. The algorithm seeks the partition of firms to clusters as well as the set of within-cluster CDFs that minimize this weighted sum of squared deviations between the empirical CDF and the candidate CDF (evaluated at the specified quantiles). In practice, we evaluate the CDF at 20 equally-spaced quantiles, and repeat the algorithm at 100 random starting values, choosing the partition associated with the starting value that achieves the lowest value of the objective function.

Expected firm premiums with firm-worker interactions. Given the firm types and the estimates of (x_i, ψ_j, θ_j), we can estimate the expected firm premiums in the model with firm-worker interactions. Using the wage equation above, the premium for a worker of type x of being employed by a firm of type B relative to a firm of type A is

$$(\psi_B + \theta_B x) - (\psi_A + \theta_A x).$$

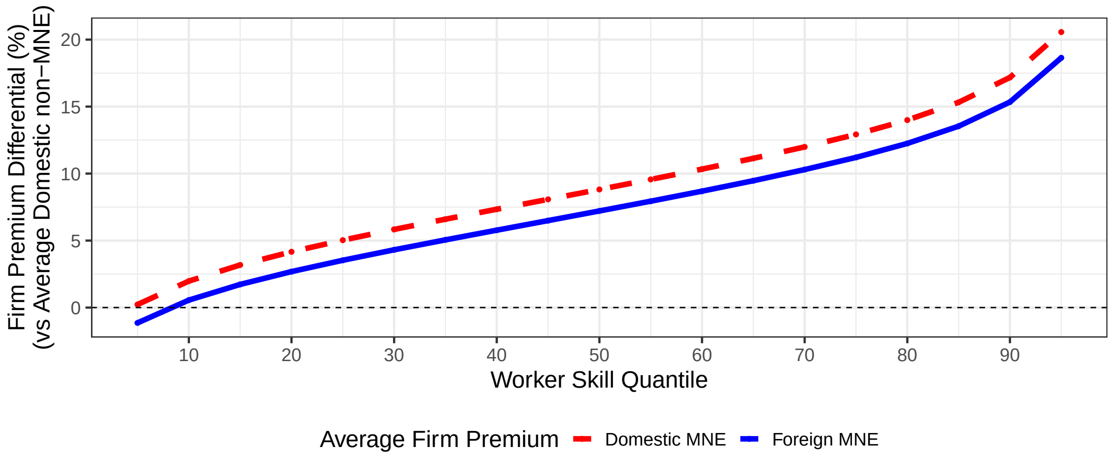
In our empirical application, we compare foreign multinationals and domestic non-multinationals. Let $P_F(k)$ denote the share of foreign multinationals that are of type k, and $P_D(k)$ denote the share of domestic non-multinationals that are of type k. For a worker of type x, the expected difference in wages when employed by a foreign firm (drawn randomly with probability $P_F(k)$) versus a domestic firm (drawn randomly with probability $P_D(k)$) is,

$$\sum_k (\psi_k + \theta_k x)(P_F(k) - P_D(k))$$

This is the expected direct effect, or foreign firm premium, for a worker of type x — it is the difference in log wages that a worker of type x is expected to receive at a randomly drawn foreign multinational versus a randomly drawn domestic non-multinational. We now estimate this quantity for various quantiles in the empirical distribution of x.

Results. Figure A.5 presents the mean difference in firm premiums between foreign and domestic firms for workers who have above average and below average quality using the estimated parameters from equation (11), finding substantial differences. We find that the foreign firm premium is monotonically increasing in the skill of workers compared to the premium offered by domestic non-multinationals to workers of the same skill. Foreign multinationals pay a 19 percent greater premium to workers in the top skill decile, but a 1 percent negative premium to workers in the bottom skill decile. Furthermore, we find that domestic-owned multinationals pay a 21 percent greater premium to workers in the top skill decile than domestic non-multinationals, but no premium to workers in the bottom skill decile. These results are consistent with multinationals having more skill-augmenting technology

Figure A.5: Firm Premiums with Firm-Worker Interactions



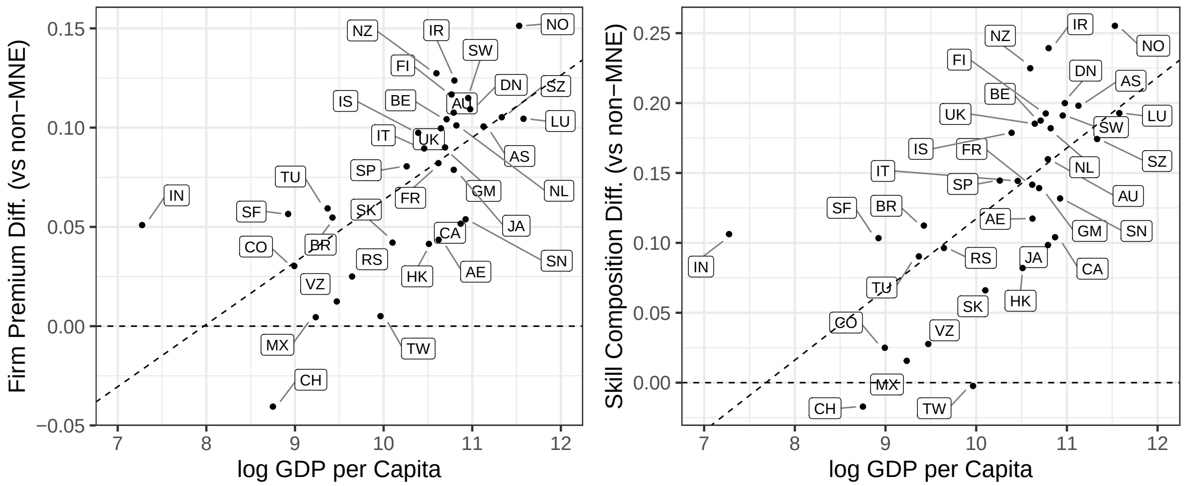
Notes: This figure presents estimates of the model in equation (11) from the grouped fixed effect estimator during 2010-2015. The horizontal axis is a quantile in the distribution of estimated worker skill level. The vertical axis is the difference in the average firm premium for a worker of a given skill level for foreign (blue) or domestic (red) multinationals, relative to the average domestic non-multinational.

than non-multinationals. Skill-augmenting technology would lead multinational firms (both foreign owned and domestic owned) to bid up the price of local labor for skilled workers such as managers, as found by Bloom et al. (2019), but not bid up the price of routine labor.

A.7 Direct Effects: Supplementary Results

Figure A.6: Direct Effects: Comparison to GDP per Capita

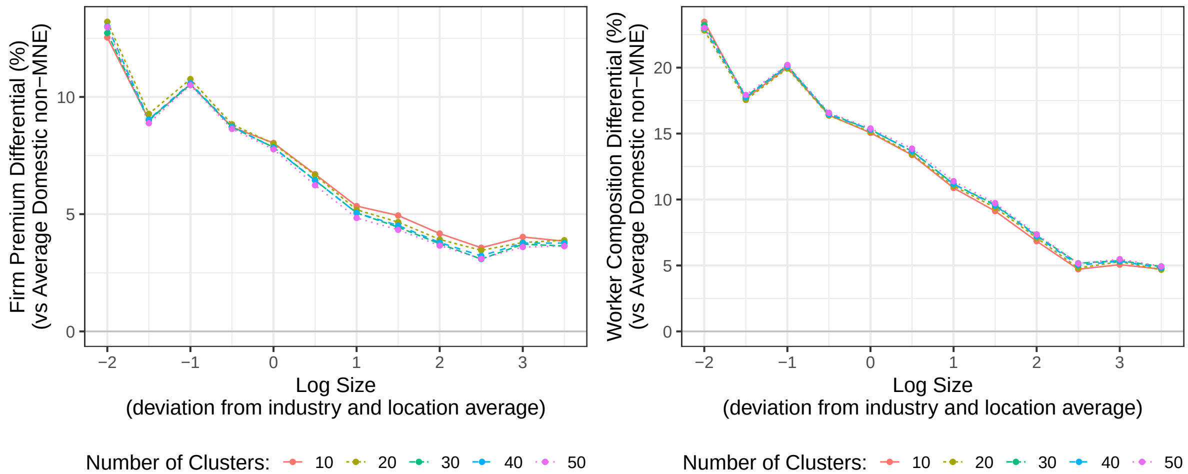
(a) Firm Premium Differential (%) (b) Skill Composition Differential (%)



Notes: This figure presents estimates of equation (10) from the grouped fixed effect estimator during 2010-2015. The vertical axis is the difference in the average firm premium (subfigure a) or average worker skill level (subfigure b) for foreign multinationals with the countries of ownership indicated by the labels, relative to the average domestic non-multinational.

Figure A.7: Direct Effects: Robustness to Number of Clusters

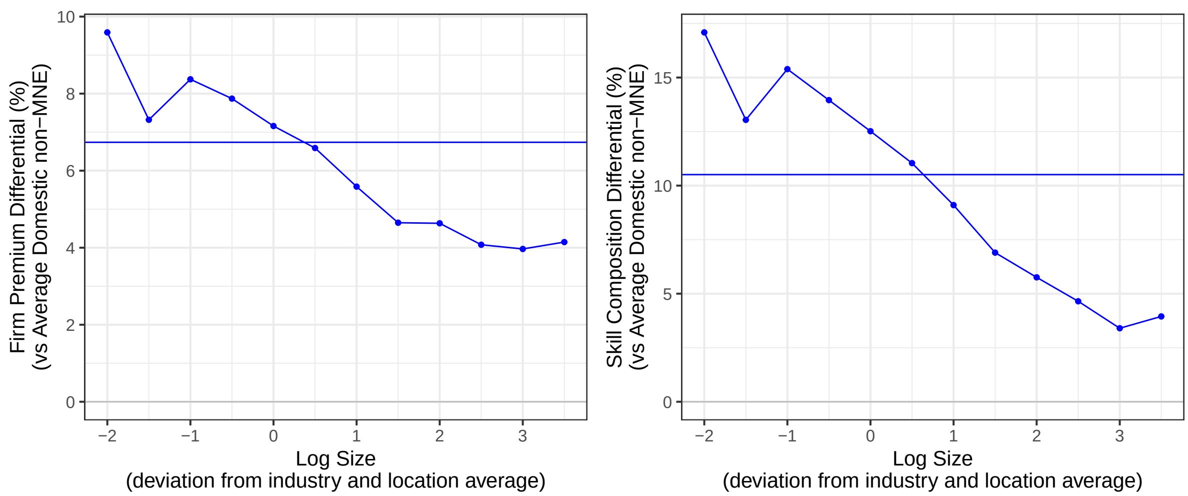
(a) Firm Premium Differential (%) (b) Skill Composition Differential (%)



Notes: This figure presents estimates of equation (10) from the grouped fixed effect estimator during 2010-2015 for different numbers of firm clusters. The horizontal axis is an equally spaced grid of width 0.5 in the residual log firm size distribution, where each unit is associated with the nearest grid point. The vertical axis is the difference in the average firm premium (subfigure a) or average worker skill level (subfigure b) for foreign multinationals, relative to the average domestic non-multinational in the same size bin.

Figure A.8: Direct Effects: Estimates for 2001-2006

(a) Firm Premium Differential (%) (b) Skill Composition Differential (%)



Notes: This figure presents estimates of the model in equation (10) from the grouped fixed effect estimator during 2001-2006. The horizontal axis is an equally spaced grid of width 0.5 in the residual log firm size distribution, where each unit is associated with the nearest grid point. The vertical axis is the difference in the average firm premium (subfigure a) or average worker skill level (subfigure b) for foreign multinationals, relative to the average domestic non-multinational in the same size bin. The horizontal lines indicate the overall averages (not conditional on a size bin).

A.8 Indirect Effects: Alternative Specifications and Robustness Checks

We now provide alternative specifications and robustness checks for the indirect effects.

Placebo tests: To improve our confidence in the orthogonality of the country of origin shocks to local growth factors, we provide a placebo test. This test uses the log changes in domestic firms' value added, wage bill, employment, and earnings of continuous workers measured in the pre-period (i.e., before the exposure shares are measured) as if they were the contemporaneous outcomes. Under our orthogonality assumption, contemporaneous country of origin shocks should not predict growth in the pre-period, conditional on the control variables. The placebo test results are presented in Appendix Table A3. The estimated second-stage coefficients become small in magnitude and statistically insignificant for all of the outcomes, consistent with our identifying assumption.

Alternative control sets: Appendix Table A4 adds controls one at a time in order to examine the sensitivity of the main results to additional controls, as well as to help understand which of the controls in our baseline specification are important. Appendix Table A5 performs the same exercise but for the OLS estimates that do not use the instrumental variable. First, as predicted above, industry-year and Census-division-year controls are important, so we include these in the baseline specification. Second, we find some marginal sensitivity to adding urban concentration controls, perhaps because of the disproportionate representation of foreign multinationals in major urban areas. Third, the results are not statistically significantly different when adding commuting zone controls for educational attainment, poverty and unemployment, or farm and manufacturing concentration. Fourth, we consider interacting the commuting-zone-year domestic employment share measure with indicators for the financial crisis of 2007-2009 as well as with all 3-year intervals in the outcome sample, finding similar results though with some loss in precision.

Controlling for past country of origin shocks: One potential concern with shift-share instruments is that, if the shocks have impacts that are slowly evolving over time, then the estimated second-stage coefficient will conflate the effects of contemporaneous and past shocks, resulting in biased estimates of the effects of contemporaneous shocks. Jaeger et al. (2018) provide theoretical justification for this type of bias in the context of immigration and propose the natural correction (i.e., controlling for lagged shocks corrects for the bias induced by lagged shocks). In column (2) of Table A6, we show that our results are nearly identical when controlling for the lagged shocks, implying that our results are not confounded by slow adjustments to past shocks.[48]

Controlling for finer industry shocks: In column (3) of Table A6, we show that the results are robust to replacing the 3-digit NAICS industry-year fixed effects with fully disaggregated 6-digit NAICS industry-year fixed effects. This suggests that we have successfully controlled for all relevant industry shocks with the baseline industry-year fixed effects.

Leaving out nearby commuting zones: A potential concern is that some workers reside in one commuting zone but commute to work in a different commuting zone nearby. As a result, workers may be affected by country of origin employment growth shocks in nearby commuting zones, which is not captured in our baseline specification. To investigate the sensitivity of our results to shocks in nearby commuting zones, we consider not only leaving out the worker's own commuting zone when constructing the shocks, but also leaving out any commuting zone within a specified radius of the worker's own commuting zone. In Appendix Table A7, we consider leaving out any commuting zone within a radius of 50 miles, 100 miles, 150 miles, 200 miles, 250 miles, or 300 miles of the worker's own commuting zone. The top of the table characterizes the distributions of the number of commuting zones left out. When using a 300-mile radius, the nearest 76 commuting zones are left out on average, with at least 117 commuting zones left out for one-fourth of the observations. Despite leaving out so many commuting zones over such a long distance, we find that the results are nearly the same, indicating that cross-commuting-zone commutes do not confound our estimates. We also consider leaving out any foreign investment in the same Census division as the worker, which amounts to leaving out 77 commuting zones on average when constructing the shocks, again finding that our results are robust to this exercise.

Excluding domestic multinationals: A possible threat to identification is that aggregate employment growth from a specific country of origin may lower transportation costs for U.S. exports to that country. For example, if Germany opens a plant in South Carolina and invests in shipping lanes from Germany to South Carolina, these shipping lanes could also be used by South Carolina domestic firms to increase exports to Germany. Although export transactions are not available in our data, most U.S. exports are carried out by multinationals (Bernard et al., 2005). When restricting the domestic sample to only non-multinationals in column (5) of Table A6, we find that the estimates are unaffected, indicating that the effects are not due to transportation costs faced by domestic exporters.

[48] In practice, we follow the implementation suggested by Borusyak et al. (2020, footnote 22). In particular, we control linearly for a lagged instrument constructed using the same exposure shares as the main instrument, but measuring the aggregate employment growth by a country of origin between $t-2$ and $t-1$ instead of between $t-1$ and t. To allow for more complicated dynamics, we also verify that results are robust to simultaneously controlling for shocks between $t-2$ and $t-1$ and between $t-3$ and $t-2$.

Excluding tax havens: A potential concern is that tax havens should not be included as foreign countries of ownership in our analysis, as some firms owned in tax havens may be misclassified domestic-owned firms. Hines (2010) classifies 52 countries as tax havens. We consider excluding all 52 tax havens from the analysis as a robustness check in column (6) of Table A6.[49] We find that the indirect effect estimates are not greatly affected or become slightly stronger when excluding tax havens.

DHS transformation of the outcome variables: Our indirect effect estimates so far have been provided for continuing domestic firms. As an alternative approach, we consider the transformation of Davis et al. (1998, "DHS") rather than log changes. The advantage of this approach is that it incorporates entry and exit into the outcome measures.[50] The results are provided in column (7) of Table A6. We find that the estimated effects become stronger, which ameliorates any concern that our main effects for continuing firms arise from survival bias. On the contrary, our results indicate net entry of domestic firms due to foreign employment growth.

Table A3: Indirect Effects Estimates: Placebo Tests

	Value Added		Employment		Wage bill		Earnings of Cont. Workers	
	Main	**Placebo**	**Main**	**Placebo**	**Main**	**Placebo**	**Main**	**Placebo**
Second-Stage:								
Coefficient	**0.96**	**-0.05**	**0.53**	**-0.17**	**0.63**	**-0.09**	**0.15**	**0.04**
(Std. Error Clustered by Commuting Zone)	(0.30)	(0.22)	(0.14)	(0.12)	(0.17)	(0.16)	(0.07)	(0.09)
(Std. Error Clustered by Country of Origin)	(0.51)	(0.45)	(0.18)	(0.14)	(0.22)	(0.16)	(0.08)	(0.09)
First-Stage:								
Coefficient	0.56	0.66	0.56	0.66	0.56	0.66	0.56	0.67
(F-statistic Clustered by Commuting Zone)	(232)	(341)	(235)	(351)	(235)	(351)	(239)	(360)
(F-statistic Clustered by Country of Origin)	(42)	(27)	(44)	(27)	(44)	(27)	(44)	(26)
Number of Firms by Commuting Zones (Millions)	41.8	36.2	46.0	38.7	46.0	38.7	44.6	37.6
Number of Workers (Millions, measured at $t-1$)	416.8	402.0	477.3	441.1	477.3	441.1	369.6	336.8

Notes: The outcome sample only includes continuing domestic firms. Observations are weighted by lagged firm size. Controls are industry-year indicators, Census-division-year indicators, measures of urban concentration, and the sum of commuting zone exposure shares. Placebo outcomes are measured as changes between $t_0 - 2$ and $t_0 - 1$, where t_0 is the time period at which the exposure shares are measured.

[49] Using the inverse HHI measure proposed by Borusyak et al. (2020), the effective number of country shocks falls from 154 to 122, indicating that we drop about one-fifth of all effective country of origin shocks. We report this inverse HHI measure for all results shown in Table A6.

[50] In particular, the DHS transformation is $2\frac{Y_{j,t}-Y_{j,t-1}}{Y_{j,t}+Y_{j,t-1}}$. If $Y_{j,t} > 0$ and $Y_{j,t-1} > 0$, this transformation is approximately $\log(Y_t) - \log(Y_{j,t-1})$. Thus, any differences between our baseline results in log changes and the results from the DHS transformation are due to firms with $Y_{j,t} \leq 0$ or $Y_{j,t-1} \leq 0$, such as firms that employ no workers at either t or $t-1$. Note that we usually weight firms by the number of workers at $t-1$. Of course, the number of workers is zero at $t-1$ for new entrants. Instead, we weight firms by the average number of workers across t and $t-1$ in the regressions with DHS transformations.

Table A4: Indirect Effect Estimates: Alternative Control Sets

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Control Specification									
CZ-year domestic employment share	✓	✓	✓	✓	✓	✓	✓	✓	✓
Industry-year fixed effects	✗	✓	✓	✓	✓	✓	✓	✓	✓
Census-division-year fixed effects	✗	✗	✓	✓	✓	✓	✓	✓	✓
CZ controls:									
Urban density measures (pre-period)	✗	✗	✗	✓	✓	✓	✓	✓	✓
Educational attainment measures (pre-period)	✗	✗	✗	✗	✓	✓	✓	✓	✓
Poverty and employment measures (pre-period)	✗	✗	✗	✗	✗	✓	✓	✓	✓
Farm and manufacturing measures (pre-period)	✗	✗	✗	✗	✗	✗	✓	✓	✓
CZ-year domestic employment share × Financial Crisis	✗	✗	✗	✗	✗	✗	✗	✓	✓
CZ-year domestic employment share × All 3-year intervals	✗	✗	✗	✗	✗	✗	✗	✗	✓
Panel A.	**Outcome: Log Change in Value Added**								
Second-Stage Coefficient	**-0.26**	**0.66**	**1.01**	**0.96**	**0.93**	**0.90**	**0.90**	**0.88**	**1.13**
(Std. Error Clustered by Commuting Zone)	(0.15)	(0.27)	(0.32)	(0.30)	(0.30)	(0.30)	(0.30)	(0.31)	(0.49)
(Std. Error Clustered by Country of Origin)	(0.37)	(0.44)	(0.50)	(0.51)	(0.50)	(0.49)	(0.49)	(0.48)	(0.90)
First-Stage Coefficient	0.96	0.61	0.56	0.56	0.56	0.56	0.56	0.55	0.44
(F-statistic Clustered by Commuting Zone)	(2,890)	(232)	(232)	(232)	(233)	(232)	(231)	(224)	(100)
(F-statistic Clustered by Country of Origin)	(1,367)	(30)	(42)	(42)	(43)	(42)	(42)	(40)	(19)
Number of Firms by Commuting Zones (Millions)	41.8	41.8	41.8	41.8	41.8	41.8	41.8	41.8	41.8
Number of Workers (Millions, measured at $t-1$)	416.8	416.8	416.8	416.8	416.8	416.8	416.6	416.6	416.6
Panel B.	**Outcome: Log Change in Employment**								
Second-Stage Coefficient	**-0.24**	**0.46**	**0.60**	**0.53**	**0.50**	**0.46**	**0.46**	**0.46**	**0.66**
(Std. Error Clustered by Commuting Zone)	(0.07)	(0.15)	(0.15)	(0.14)	(0.13)	(0.13)	(0.13)	(0.13)	(0.21)
(Std. Error Clustered by Country of Origin)	(0.18)	(0.20)	(0.22)	(0.18)	(0.18)	(0.17)	(0.17)	(0.17)	(0.30)
First-Stage Coefficient	0.95	0.61	0.56	0.56	0.56	0.56	0.56	0.56	0.44
(F-statistic Clustered by Commuting Zone)	(2,890)	(233)	(235)	(235)	(236)	(235)	(234)	(227)	(102)
(F-statistic Clustered by Country of Origin)	(1,402)	(30)	(44)	(44)	(44)	(44)	(43)	(42)	(19)
Number of Firms by Commuting Zones (Millions)	46.0	46.0	46.0	46.0	46.0	46.0	45.9	45.9	45.9
Number of Workers (Millions, measured at $t-1$)	477.3	477.3	477.3	477.3	477.3	477.3	477.1	477.1	477.1
Panel C.	**Outcome: Log Change in Wage Bill**								
Second-Stage Coefficient	**-1.19**	**0.54**	**0.70**	**0.63**	**0.59**	**0.55**	**0.54**	**0.53**	**0.72**
(Std. Error Clustered by Commuting Zone)	(0.14)	(0.18)	(0.18)	(0.17)	(0.16)	(0.16)	(0.16)	(0.16)	(0.26)
(Std. Error Clustered by Country of Origin)	(0.36)	(0.25)	(0.25)	(0.22)	(0.22)	(0.21)	(0.20)	(0.20)	(0.35)
First-Stage Coefficient	0.95	0.61	0.56	0.56	0.56	0.56	0.56	0.56	0.44
(F-statistic Clustered by Commuting Zone)	(2,890)	(233)	(235)	(235)	(236)	(235)	(234)	(227)	(102)
(F-statistic Clustered by Country of Origin)	(1,402)	(30)	(44)	(44)	(44)	(44)	(43)	(42)	(19)
Number of Firms by Commuting Zones (Millions)	46.0	46.0	46.0	46.0	46.0	46.0	45.9	45.9	45.9
Number of Workers (Millions, measured at $t-1$)	477.3	477.3	477.3	477.3	477.3	477.3	477.1	477.1	477.1
Panel D.	**Outcome: Log Change in Earnings of Cont. Workers**								
Second-Stage Coefficient	**-1.21**	**0.11**	**0.17**	**0.15**	**0.13**	**0.13**	**0.13**	**0.11**	**0.13**
(Std. Error Clustered by Commuting Zone)	(0.08)	(0.08)	(0.08)	(0.07)	(0.07)	(0.07)	(0.07)	(0.07)	(0.11)
(Std. Error Clustered by Country of Origin)	(0.23)	(0.11)	(0.09)	(0.08)	(0.07)	(0.07)	(0.07)	(0.07)	(0.13)
First-Stage Coefficient	0.95	0.61	0.56	0.56	0.56	0.56	0.56	0.56	0.44
(F-statistic Clustered by Commuting Zone)	(2,930)	(238)	(239)	(239)	(240)	(239)	(238)	(231)	(104)
(F-statistic Clustered by Country of Origin)	(1,340)	(29)	(43)	(44)	(44)	(43)	(43)	(41)	(19)
Number of Firms by Commuting Zones (Millions)	44.6	44.6	44.6	44.6	44.6	44.6	44.6	44.6	44.6
Number of Workers (Millions, measured at $t-1$)	369.6	369.6	369.6	369.6	369.6	369.6	369.5	369.5	369.5

Notes: The outcome sample only includes continuing domestic firms. Observations are weighted by lagged firm size. Controls are indicated at the top of the table. Our baseline control set is in column (4).

Table A5: Indirect Effect Estimates: OLS Estimates for Various Control Sets

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Control Specification									
CZ-year domestic employment share	✓	✓	✓	✓	✓	✓	✓	✓	✓
Industry-year fixed effects	✗	✓	✓	✓	✓	✓	✓	✓	✓
Census-division-year fixed effects	✗	✗	✓	✓	✓	✓	✓	✓	✓
CZ controls:									
Urban density measures (pre-period)	✗	✗	✗	✓	✓	✓	✓	✓	✓
Educational attainment measures (pre-period)	✗	✗	✗	✗	✓	✓	✓	✓	✓
Poverty and employment measures (pre-period)	✗	✗	✗	✗	✗	✓	✓	✓	✓
Farm and manufacturing measures (pre-period)	✗	✗	✗	✗	✗	✗	✓	✓	✓
CZ-year domestic employment share × Financial Crisis	✗	✗	✗	✗	✗	✗	✗	✓	✓
CZ-year domestic employment share × All 3-year intervals	✗	✗	✗	✗	✗	✗	✗	✗	✓
Panel A.	**Outcome: Log Change in Value Added**								
OLS Coefficient	**0.05**	**0.49**	**0.34**	**0.33**	**0.33**	**0.32**	**0.32**	**0.31**	**0.31**
(Std. Error Clustered by Commuting Zone)	(0.13)	(0.10)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)
Number of Firms by Commuting Zones (Millions)	41.8	41.8	41.8	41.8	41.8	41.8	41.8	41.8	41.8
Number of Workers (Millions, measured at $t-1$)	416.8	416.8	416.8	416.8	416.8	416.8	416.6	416.6	416.6
Panel B.	**Outcome: Log Change in Employment**								
OLS Coefficient	**-0.03**	**0.28**	**0.24**	**0.22**	**0.22**	**0.21**	**0.21**	**0.21**	**0.22**
(Std. Error Clustered by Commuting Zone)	(0.07)	(0.05)	(0.05)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)
Number of Firms by Commuting Zones (Millions)	46.0	46.0	46.0	46.0	46.0	46.0	45.9	45.9	45.9
Number of Workers (Millions, measured at $t-1$)	477.3	477.3	477.3	477.3	477.3	477.3	477.1	477.1	477.1
Panel C.	**Outcome: Log Change in Wage Bill**								
OLS Coefficient	**-0.58**	**0.39**	**0.31**	**0.29**	**0.30**	**0.29**	**0.29**	**0.28**	**0.29**
(Std. Error Clustered by Commuting Zone)	(0.13)	(0.07)	(0.05)	(0.05)	(0.05)	(0.05)	(0.05)	(0.05)	(0.05)
Number of Firms by Commuting Zones (Millions)	46.0	46.0	46.0	46.0	46.0	46.0	45.9	45.9	45.9
Number of Workers (Millions, measured at $t-1$)	477.3	477.3	477.3	477.3	477.3	477.3	477.1	477.1	477.1
Panel D.	**Outcome: Log Change in Earnings of Cont. Workers**								
OLS Coefficient	**-0.68**	**0.16**	**0.13**	**0.13**	**0.13**	**0.13**	**0.13**	**0.12**	**0.13**
(Std. Error Clustered by Commuting Zone)	(0.09)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
Number of Firms by Commuting Zones (Millions)	44.6	44.6	44.6	44.6	44.6	44.6	44.6	44.6	44.6
Number of Workers (Millions, measured at $t-1$)	369.6	369.6	369.6	369.6	369.6	369.6	369.5	369.5	369.5

Notes: The outcome sample only includes continuing domestic firms. Observations are weighted by lagged firm size. Controls are indicated at the top of the table. Our baseline control set is in column (4).

Table A6: Indirect Effects Estimates: Robustness Checks

	(1) Baseline	(2) Control Lag IV	(3) Control NAICS-6	(4) Leave out 300m Radius	(5) Exclude Dom MNE	(6) Exclude Tax Havens	(7) DHS Transform
Panel A.	**Outcome: Log Change in Value Added**						
Second-Stage Coefficient	**0.96**	**0.98**	**0.93**	**0.97**	**0.87**	**1.10**	**1.44**
(Std. Error Clustered by Commuting Zone)	(0.30)	(0.32)	(0.26)	(0.35)	(0.25)	(0.36)	(0.47)
(Std. Error Clustered by Country of Origin)	(0.51)	(0.51)	(0.34)	(0.51)	(0.35)	(0.62)	(0.53)
First-Stage Coefficient	0.56	0.54	0.56	0.59	0.57	0.55	0.57
(F-statistic Clustered by Commuting Zone)	(232)	(208)	(233)	(143)	(260)	(233)	(264)
(F-statistic Clustered by Country of Origin)	(42)	(42)	(43)	(42)	(40)	(35)	(41)
Number of Firms by Commuting Zones (Millions)	41.8	40.5	41.8	41.8	40.6	41.8	66.6
Number of Workers (Millions, measured at $t-1$)	416.8	401.0	416.8	416.8	344.1	416.8	497.8
Effective Number of Country Shocks (Inverse HHI)	153.6	153.6	153.6	153.6	153.6	122.2	153.6
Panel B.	**Outcome: Log Change in Employment**						
Second-Stage Coefficient	**0.53**	**0.56**	**0.50**	**0.56**	**0.42**	**0.55**	**0.83**
(Std. Error Clustered by Commuting Zone)	(0.14)	(0.15)	(0.14)	(0.17)	(0.14)	(0.17)	(0.29)
(Std. Error Clustered by Country of Origin)	(0.18)	(0.18)	(0.17)	(0.18)	(0.18)	(0.22)	(0.23)
First-Stage Coefficient	0.56	0.54	0.56	0.59	0.57	0.55	0.58
(F-statistic Clustered by Commuting Zone)	(235)	(211)	(236)	(145)	(258)	(241)	(270)
(F-statistic Clustered by Country of Origin)	(44)	(43)	(44)	(44)	(42)	(37)	(41)
Number of Firms by Commuting Zones (Millions)	46.0	44.6	46.0	46.0	44.6	46.0	69.2
Number of Workers (Millions, measured at $t-1$)	477.3	459.6	477.3	477.3	395.6	477.3	519.7
Effective Number of Country Shocks (Inverse HHI)	153.6	153.6	153.6	153.6	153.6	122.2	153.6
Panel C.	**Outcome: Log Change in Wage Bill**						
Second-Stage Coefficient	**0.63**	**0.61**	**0.59**	**0.69**	**0.49**	**0.67**	**0.88**
(Std. Error Clustered by Commuting Zone)	(0.17)	(0.19)	(0.17)	(0.21)	(0.17)	(0.20)	(0.29)
(Std. Error Clustered by Country of Origin)	(0.22)	(0.22)	(0.21)	(0.22)	(0.22)	(0.27)	(0.26)
First-Stage Coefficient	0.56	0.54	0.56	0.59	0.57	0.55	0.58
(F-statistic Clustered by Commuting Zone)	(235)	(211)	(236)	(145)	(258)	(241)	(270)
(F-statistic Clustered by Country of Origin)	(44)	(43)	(44)	(44)	(42)	(37)	(41)
Number of Firms by Commuting Zones (Millions)	46.0	44.6	46.0	46.0	44.6	46.0	69.2
Number of Workers (Millions, measured at $t-1$)	477.3	459.6	477.3	477.3	395.6	477.3	519.7
Effective Number of Country Shocks (Inverse HHI)	153.6	153.6	153.6	153.6	153.6	122.2	153.6
Panel D.	**Outcome: Log Change in Earnings of Continuing Workers**						
Second-Stage Coefficient	**0.15**	**0.10**	**0.13**	**0.17**	**0.09**	**0.15**	**0.15**
(Std. Error Clustered by Commuting Zone)	(0.07)	(0.08)	(0.07)	(0.09)	(0.07)	(0.08)	(0.07)
(Std. Error Clustered by Country of Origin)	(0.08)	(0.08)	(0.07)	(0.08)	(0.07)	(0.09)	(0.08)
First-Stage Coefficient	0.56	0.55	0.56	0.59	0.57	0.56	0.56
(F-statistic Clustered by Commuting Zone)	(239)	(214)	(240)	(150)	(265)	(247)	(239)
(F-statistic Clustered by Country of Origin)	(44)	(43)	(44)	(44)	(41)	(37)	(44)
Number of Firms by Commuting Zones (Millions)	44.6	43.3	44.6	44.6	43.3	44.6	44.6
Number of Workers (Millions, measured at $t-1$)	369.6	356.0	369.6	369.6	304.3	369.6	369.6
Effective Number of Country Shocks (Inverse HHI)	153.6	153.6	153.6	153.6	153.6	122.2	153.6

Notes: The outcome sample only includes continuing domestic firms (unless otherwise specified). Observations are weighted by lagged firm size (unless otherwise specified). Controls are industry-year indicators, Census-division-year indicators, measures of urban concentration, and the sum of commuting zone exposure shares (unless otherwise specified).

Table A7: Indirect Effects Estimates: Leave-out Specifications

	Leave out No CZ (include Own)	**Leave out Own CZ**	**Leave out CZs within Radius (based on nearest distance in miles)**						**Leave out Entire Census Division**
			50	100	150	200	250	300	
			Number of CZs excluded						
Mean	0	1	7	16	28	42	59	76	77
25th quantile	0	1	5	10	17	24	30	37	58
50th quantile	0	1	8	16	28	42	57	74	84
75th quantile	0	1	9	21	38	59	84	114	104
Panel A.	**Outcome: Log Change in Value Added**								
Second-Stage Coefficient	**0.87**	**0.96**	**0.96**	**0.94**	**0.95**	**0.97**	**1.01**	**0.97**	**0.89**
(Std. Error Clustered by Commuting Zone)	(0.28)	(0.30)	(0.32)	(0.33)	(0.32)	(0.33)	(0.33)	(0.35)	(0.32)
(Std. Error Clustered by Country of Origin)	(0.51)	(0.51)	(0.51)	(0.51)	(0.51)	(0.51)	(0.51)	(0.51)	(0.51)
First-Stage Coefficient	0.59	0.56	0.56	0.56	0.58	0.58	0.59	0.59	0.61
(F-statistic Clustered by Commuting Zone)	(273)	(232)	(212)	(199)	(185)	(172)	(161)	(143)	(189)
(F-statistic Clustered by Country of Origin)	(42)	(42)	(42)	(42)	(42)	(42)	(42)	(42)	(42)
Number of Firms by Commuting Zones (Millions)	41.8	41.8	41.8	41.8	41.8	41.8	41.8	41.8	41.8
Number of Workers (Millions, measured at $t-1$)	416.8	416.8	416.8	416.8	416.8	416.8	416.8	416.8	416.8
Panel B.	**Outcome: Log Change in Employment**								
Second-Stage Coefficient	**0.52**	**0.53**	**0.54**	**0.55**	**0.55**	**0.55**	**0.57**	**0.56**	**0.55**
(Std. Error Clustered by Commuting Zone)	(0.13)	(0.14)	(0.15)	(0.15)	(0.15)	(0.16)	(0.16)	(0.17)	(0.15)
(Std. Error Clustered by Country of Origin)	(0.18)	(0.18)	(0.18)	(0.18)	(0.18)	(0.18)	(0.18)	(0.18)	(0.18)
First-Stage Coefficient	0.59	0.56	0.56	0.56	0.58	0.58	0.59	0.59	0.61
(F-statistic Clustered by Commuting Zone)	(277)	(235)	(214)	(201)	(188)	(175)	(163)	(145)	(192)
(F-statistic Clustered by Country of Origin)	(44)	(44)	(44)	(44)	(44)	(44)	(44)	(44)	(44)
Number of Firms by Commuting Zones (Millions)	46.0	46.0	46.0	46.0	46.0	46.0	46.0	46.0	46.0
Number of Workers (Millions, measured at $t-1$)	477.3	477.3	477.3	477.3	477.3	477.3	477.3	477.3	477.3
Panel C.	**Outcome: Log Change in Wage Bill**								
Second-Stage Coefficient	**0.61**	**0.63**	**0.64**	**0.64**	**0.65**	**0.66**	**0.69**	**0.69**	**0.64**
(Std. Error Clustered by Commuting Zone)	(0.16)	(0.17)	(0.18)	(0.18)	(0.19)	(0.19)	(0.20)	(0.21)	(0.18)
(Std. Error Clustered by Country of Origin)	(0.22)	(0.22)	(0.22)	(0.22)	(0.22)	(0.22)	(0.22)	(0.22)	(0.22)
First-Stage Coefficient	0.59	0.56	0.56	0.56	0.58	0.58	0.59	0.59	0.61
(F-statistic Clustered by Commuting Zone)	(277)	(235)	(214)	(201)	(188)	(175)	(163)	(145)	(192)
(F-statistic Clustered by Country of Origin)	(44)	(44)	(44)	(44)	(44)	(44)	(44)	(44)	(44)
Number of Firms by Commuting Zones (Millions)	46.0	46.0	46.0	46.0	46.0	46.0	46.0	46.0	46.0
Number of Workers (Millions, measured at $t-1$)	477.3	477.3	477.3	477.3	477.3	477.3	477.3	477.3	477.3
Panel D.	**Outcome: Log Change in Earnings of Continuing Workers**								
Second-Stage Coefficient	**0.15**	**0.15**	**0.14**	**0.14**	**0.15**	**0.16**	**0.17**	**0.17**	**0.14**
(Std. Error Clustered by Commuting Zone)	(0.07)	(0.07)	(0.07)	(0.08)	(0.08)	(0.08)	(0.08)	(0.09)	(0.08)
(Std. Error Clustered by Country of Origin)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)
First-Stage Coefficient	0.59	0.56	0.56	0.56	0.58	0.59	0.59	0.59	0.61
(F-statistic Clustered by Commuting Zone)	(281)	(239)	(218)	(205)	(192)	(179)	(167)	(150)	(196)
(F-statistic Clustered by Country of Origin)	(44)	(44)	(44)	(44)	(44)	(44)	(44)	(44)	(44)
Number of Firms by Commuting Zones (Millions)	44.6	44.6	44.6	44.6	44.6	44.6	44.6	44.6	44.6
Number of Workers (Millions, measured at $t-1$)	369.6	369.6	369.6	369.6	369.6	369.6	369.6	369.6	369.6

Notes: The outcome sample only includes continuing domestic firms. Observations are weighted by lagged firm size. Controls are industry-year indicators, Census-division-year indicators, measures of urban concentration, and the sum of commuting zone exposure shares. Our baseline specification is in column (2).