U.S. MULTINATIONALS AND PREFERENTIAL MARKET ACCESS

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Abstract

We combine firm level panel data on U.S. foreign affiliate activity with detailed measures of U.S. trade policy to study the relationship between offshoring and preferential market access. Consistent with theory, we find that trade preferences and offshoring activity are positively and significantly correlated. Using instrumental variables, we estimate that a 10% increase in U.S. foreign affiliate exports to the U.S. is associated with a 4 percentage point increase in the rate of preferential duty-free access. Restricting attention to the Generalized System of Preferences (GSP) among developing countries, this estimate more than triples relative to the baseline, full sample results.

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1. Overview

Recent theoretical work in Blanchard (2007, 2010) suggests that the pattern of international investment and multinational enterprise (MNE) activity may play an important role in shaping government preferences over trade policies: when a multinational firm owns export-oriented (i.e. 'offshoring') affiliates abroad, the MNE's 'home' country government has an incentive to improve market access for imports from those foreign affiliates, for the simple reason that greater market access means higher rates of return to the government's MNE constituents. To the extent that governments respond to the interests of constituent industries, differences in the pattern of firm operations across the globe may be reflected in trade policy.

Unfortunately, this clear empirical prediction is not easily taken to the data. Empirically testing the hypothesis that MNE operations and foreign direct investment (FDI) influence governments' most preferred trade policies proves problematic in three key dimensions. First, most advanced economies have set tariffs cooperatively since the inception of the General Agreement on Tariffs and Trade (GATT) in 1947. Since tariff concessions are negotiated multilaterally and are subject to the Most Favored Nation (MFN) non-discrimination clause, the econometrician is challenged to distinguish the influence of FDI from other multilateral pressures at the negotiating table. A government may be unwilling to change its MFN tariff on a particular good if its underlying objective is to improve market access for just a handful of countries. Moreover, many of the MFN tariff concessions negotiated within the GATT/WTO framework apply to broad classes of goods rather than narrowly defined HTS-8 categories, further diluting the influence that country- and product-specific FDI and MNE activity could have on MFN tariffs.

Our empirical strategy sidesteps this problem by focusing on the potential influence of MNE activity not on MFN tariffs, but on preferential treatment of trade flows. Focusing our analysis on U.S. trade policy, we study the pattern of duty-free access offered under special trade initiatives (e.g. AGOA), free trade agreements (FTAs), and the Generalized System of Preferences (GSP), the 'aid-through-trade' initiative in which industrial nations grant expanded market access to developing countries. Because preferential treatment is exempt from MFN, trade preferences may offer a closer reflection of governments' unilateral trade policy preferences.[1]

A second potential complication for empirical testing lies in differentiating export-oriented (vertical) FDI apart from market-seeking (horizontal) FDI. While theory predicts that export-oriented FDI will lower tariffs in the investment-source country, the same is not true of market-seeking investment. If multinationals operate horizontal 'tariff jumping' operations abroad, those activities will have either no effect, or – via general equilibrium forces – a small *positive* effect on the investment-source country's optimal tariff. Fortunately, rich data from the U.S. Bureau of Economic Analysis (BEA) offer an empirical solution. Based on detailed surveys of U.S. multinational firms, we are able to measure foreign affiliate sales by destination, which allows us to distinguish export-oriented affiliate activity (measured as U.S. MNE subsidiaries' sales of goods back to the U.S.) apart from horizontal activity (subsidiaries' sales to the foreign local market).

The third and most challenging hurdle in identifying a potential effect of MNE activity on trade policy is the presumed endogeneity of export platform investment. We expect that MNEs will be more likely to set up offshoring operations in countries that have preferential market access to the U.S. The theory in Blanchard (2007) makes this reverse causality explicit, demonstrating that export-oriented FDI will in general increase as tariffs to the export-destination market are lowered.

Our response is to use an instrumental variables (IV) approach to control for the endogeneity of export-oriented FDI. The theory makes it clear that absent general equilibrium forces, horizontal investment should not itself influence or be influenced by U.S. tariffs.[2] In practice, however, we expect market-seeking FDI to be positively correlated with export-oriented FDI, since both rely on a favorable climate for investment. Thus, we use (horizontal) MNE sales to the local market as an instrument for (export-oriented) MNE goods sales to the U.S. Crucially, we construct our instrument using sales by *only* those multinational affiliates that do not also sell goods back to the U.S., which addresses the potential concern

[1]MFN exemptions are remarkably widespread in practice. See Section 2 for discussion of the institutional structure governing U.S. trade policy and Figure 2 for the extent of variation in our data. FTAs are exempt from MFN under GATT Article XXIV, while the Enabling Clause exempts the GSP.

[2]In the context of our country-industry-level study, we do not expect general equilibrium effects to manifest in practice and we find no evidence that they do. (See, e.g., column 5 in Table 4.)

that increasing returns to scale at the affiliate level could lead to violations of the exclusion restriction.[3]

Equipped with the empirical strategy described above, we assemble a three-way panel data set including 80 industries, 184 countries, and ten years (1997-2006) to measure the responsiveness of U.S. trade preferences to export-oriented, offshoring activities by U.S. MNEs. Our findings are consistent with a causal relationship: conservatively, a 10% increase in MNE affiliates' exports to the U.S. is associated with an increase in the duty-free access of about 4 percentage points (roughly 20% relative to the mean value of preferential market access), controlling for the endogeneity of MNE foreign affiliates' exports to the U.S. Among potentially GSP-eligible developing countries, the estimated effect is more than three times higher: a 10% increase in affiliate goods exports to the U.S. is associated with a roughly 15 percentage point increase in the rate of duty-free access under the GSP. These baseline estimates prove robust in a host of alternative empirical specifications.

Our results offer compelling evidence that offshoring MNE activity can spur preferential trade liberalization to the MNE's home country, which in turn may deepen economic integration between the investment host and source countries even further, but may also lead to substantial trade and investment diversion. A particular concern is that while some trading partners experience ever-greater economic integration through this investment-trade complementarity, other countries may be left out entirely.

This paper complements and considerably extends the empirical literature on the determinants of preferential treatment and the influence of international investment.[4] Most relevant to this paper's objective are Magee (2003) and Baier and Bergstrand (2004), which empirically test the determinants of preferential trade agreements (PTAs). Both papers view preferential agreements as a binary, country-level decision and do not exploit, as we do, the considerable variation in preferential treatment at the industry level. Also related are DeVault (1996), Özden and Reinhardt (2005), Lederman and Özden (2007), Kee, Olarreaga, and Silva (2007), and Ludema, Mayda, and Mishra (2011), which explore the domestic and

[3] One might worry that U.S. trade policy could influence an affiliate's horizontal FDI sales if the decision to enter (or expand in) the market depends on both the level of exports back to the U.S. and local sales, as could arise under increasing returns to scale.

[4] See Bhagwati, Krishna, and Panagariya (1999) for a broad review.

geo-political determinants of U.S. trade preferences and (in the last paper) U.S. tariff suspensions. None of these studies consider a role for foreign direct investment or multinational firms.

A second subset of related work examines the role of (exogenous) PTAs in determining investment flows. The seminal theoretical work by Motta and Norman (1996) identifies the potential for PTAs to attract both export-oriented and import-competing investment. Treating PTAs as exogenous, Balasubramanyam, Sapsford, and Griffiths (2002) find no effect of regional integration agreements on bilateral FDI flows in a cross section of countries in 1995. To our knowledge, this is the only study apart from our own to evaluate empirically the relationship between investment flows and trade agreements, though it examines the reverse causality and does not address the question of endogeneity.

Finally this paper complements a companion project, Blanchard and Matschke (2006), in which we extend a variant of Magee (2003) and Baier and Bergstrand (2004) to study the effect of bilateral investment positions on preferential trade agreements. A key finding from that exercise is the importance of using industry-level investment data to estimate the effects of export-oriented investment apart from the influence of market-seeking investment, as we do here.

The remainder of the paper proceeds as follows. Section 2 reviews the theoretical motivation behind our empirical approach and discusses the political process through which U.S. trade policy solicits and responds to multinational firms' concerns. Section 3 outlines our empirical strategy; Section 4 describes the data; and Section 5 presents the results. We describe a series of extensions and robustness tests in Section 6 before concluding in Section 7.

2. Theoretical Motivation and Political Mechanisms

In a general equilibrium framework, Blanchard (2010) demonstrates that export-oriented foreign investment reduces the optimal (bilateral) tariff imposed by the investment source country. The mechanism is straightforward. When a country is large, its import tariffs shift part of the tax burden onto foreign exporters, generating the well-known 'terms of trade cost shifting' externality.[5] When part of the foreign export sector is owned by domestic constituents in the importing country (for instance, through offshoring investments by MNEs), the terms of trade externality is internalized, causing the optimal tariff to decline.[6] If trade protection is politically motivated and offshore investors actively lobby their local government, tariffs should fall even further. In a many-country world, Blanchard (2007) shows that a large importing country has an incentive to use MFN exemptions to favor its overseas investors rather than change MFN tariffs, particularly if the investment host countries are small.

Our most important empirical prediction follows directly: all else equal, offshoring activity by firms from a given source country – here the U.S.– in a particular host country and industry will increase the incentives for U.S. policy makers to improve market access for imports from U.S.-affiliated producers overseas.[7] Because trade policy by law cannot discriminate at the firm level (except in anti-dumping cases), we expect this mechanism to operate at the (HTS-8) tariff line level on a country-to-country basis.

The degree to which trade preferences will respond to MNE offshoring operations depends on four key factors: the sensitivity of MNE investors' profits to U.S. trade policy; the political influence of the MNEs relative to import-competing domestic groups; the opportunity cost of lost tariff revenues; and the extent to which preferences offered to one country and industry dilute the MNE profits derived from rest of the world (ROW) imports.[8] Translating to empirically testable predictions in the context of U.S. trade preferences and multinational affiliate activity, we thus draw the following predictions:

Empirical Prediction 1. *U.S. trade preferences for product j from country c will be:*

(*i*) *increasing in U.S. MNE affiliates' sales of product j from country c to the U.S.;*

(*ii*) *decreasing in U.S. domestic sales of product j;*

(*iii*) *decreasing in total U.S. imports of product j from country c;*

(*iv*) *decreasing in total U.S. imports of product j from U.S. MNE affiliates in the ROW.*

[5] See Bagwell and Staiger (1999) for the canonical application to GATT/WTO rules, and Johnson (1951-52) for the foundational exposition of the optimal tariff argument.

[6] While offshoring operations are only one form of export-oriented foreign investment, they are also the most easily measurable, which is why we use them as the focus of our study.

[7] See Lemma 3.5 in Blanchard (2007) for formal treatment.

[8] The first three predictions can be found in both Blanchard (2007) or (2010), while the last derives from the multi-country framework in Blanchard (2007) for the case of a finite number of trading partners.

A few points of explanation bear mention. First, predictions (*i*) and (*iv*) use MNE affiliate sales of goods back to the U.S. as the relevant measure of the sensitivity of U.S. multinational investors' profits to U.S. trade policy; our choice of this measure rests on the assumption that an increase in preferential market access will translate into higher local per-unit revenues and profits earned by those multinational affiliates supplying the U.S.[9] In prediction (*ii*), we use the size of domestic U.S. sales as a proxy for potential protectionist pressure, with greater domestic sales associated with a more protectionist stance. In our baseline empirical specifications, we also include a quadratic term for U.S. domestic sales to capture potential concavity in the political response, loosely in the spirit of Bombardini and Trebbi (2011).[10] Prediction (*iii*) reflects the opportunity cost of tariff preferences, which is greater the larger the volume of imports that would no longer pay MFN duties. Prediction (*iv*) captures the possibility that enhanced market access for one country could, via trade diversion, cannibalize profits from other MNE affiliate suppliers to the U.S. in the rest of the world.

In addition, we expect that both MFN tariffs and total U.S. industry imports could be important, but the sign prediction is ambiguous. A higher MFN tariff implies a greater opportunity cost of offering tariff preferences, but at the same time, greater MFN tariffs may induce more active MNE political lobbying for discretionary policies. The net influence of the level of total U.S. imports is similarly indeterminate: all else equal, greater total U.S. imports may imply less domestic political pressure against preferences. Conversely, higher total U.S. imports imply a greater tariff revenue loss from the rest of the world via trade diversion.

Political Mechanisms. We now offer a quick overview of the political process that determines U.S. trade preferences. We highlight three important points that both support our empirical approach and elucidate the underlying mechanism. Our first point is simply that we may treat MFN tariffs as pre-determined in the context of our empirical analysis. U.S. MFN tariff bindings were last set in 1994 during the Uruguay Round negotiations –

[9]By Hotelling's lemma, recall that the derivative of the profit function with respect to output price p is supply (i.e. $\frac{\partial \pi(p,\vec{w})}{\partial p} = y(p,\vec{w})$). Thus, if foreign investors are the residual claimants of affiliate profits and view local wages as fixed, then our sales measure is *exactly* the derivative that theory suggests.

[10]In section 6, we also consider explicit measures of political activity by domestic and MNE actors using indicators from Gawande and Bandyopadhyay (2000) and Bombardini (2008).

three years before the start of our sample period in 1997. Applied tariffs have moved little since, and the variation that has occurred has been the result of extended phase-in schedules negotiated before the beginning of our sample period (and thus independent of conditions after 1997).

Second, U.S. trade rules provide considerable scope for discretionary application of trade preferences in practice.[11] While the MFN clause of the GATT limits the extent to which WTO members can offer other members preferential market access, there are widespread exceptions. FTAs offer preferential market access at the bilateral level and include sharp carve-outs and differential phase-in periods at the product level that introduce the potential for discretionary application across goods in practice. Likewise, the GSP, meant to allow countries to offer preferential market access to developing countries, introduces the potential for discretion at the country-product level.[12] Under the GSP, so-called 'competitive need limitations' waivers and exemptions, differentially binding rules of origin, and product-level application of country penalties lead to substantial variation in applied preferences at the country-product-year level.[13]

The final point spotlights the underlying mechanism: we directly observe multinational firms publicly and actively engaged in the political process that determines the structure of U.S. trade preferences.[14] For example, during a 2005-2006 USTR review of country-level GSP eligibility, Dana Corporation, a large U.S. vehicle parts manufacturer, requested extension of GSP preferences for its offshore operations in Argentina, Brazil, India, and Venezuela, writing, "elimination of GSP benefits...will result in significant harm to Dana's foreign investments..."[15] During the same year, Alcoa, a U.S. aluminum producer, petitioned for continued GSP eligibility for Brazil, Russia, and Venezuela. In its written comments to the USTR, Alcoa argued that loss of GSP access would "cause significant disruption to [its]

[11] Article XXIV of the GATT defines the rules for PTAs; for the U.S. GSP, duty-free treatment is specified by Title V of the U.S. Trade Act of 1974 (19 U.S.C. 2461).

[12] See Blanchard and Hakobyan (2014) for a detailed analysis of discretion under the U.S. GSP.

[13] For example, following a 1997 violation of intellectual property rights in pharmaceuticals, Argentina lost GSP access for a variety of products ranging from anchovies to raw cane sugar (ibid., footnote 2). Our baseline preference measure captures any differential uptake in GSP preferences (whether by formal eligibility or by exporter choice); our alternative eligibility measure reflects competitive need limitations and other formal exclusions, but cannot capture differentially binding rules of origin.

[14] These examples come from public petitions to the U.S. Trade Representative (USTR) and bear no relationship to the BEA data. Additional examples and documentation available per request.

[15] 2006 GSP Annual Review, written comment of Dana Corporation (USTR docket FR-0052)(14)

supply chain" and could cause "the imposition of over $3 million in additional costs."[16] It went on to add, "If Brazil, Russia, and Venezuela have their eligibility restricted in some way, we request that the following products not be removed," before listing the relevant HTS codes for its imports from those countries. There are many such examples, and they occur every year.[17]

With a clear understanding of the theoretical motivation and the institutional mechanisms that drive trade preferences in practice, we now turn to the empirical analysis.

3. Empirical Strategy

Let the latent variable θ^*_{cjt} represent the underlying U.S. preference for offering duty-free market access to imports of a given product j from country c in year t.[18] Under U.S. trade law, observed preferences θ_{cjt} must be binary: imports of a given product j from country c are either eligible for duty-free access or subject to the MFN rate.[19] Because several key variables (including our measure for multinational activity) are only available at the industry level, however,[20] we need to aggregate our dependent variable to the 4-digit North American Industry Classification System (NAICS-4) level. Indexing industries by i, our observed preference measure $\theta_{cit} \in [0, 1]$ denotes the share of imports from country c in industry i that are eligible for duty free access to the U.S. in year t.

We define the econometric model:

$$\theta_{cit} = \alpha_0 + \alpha_1 FDI_{cit} + \beta \cdot X_{cit} + \gamma_c + \delta_i + \eta_t + \epsilon_{cit}, \tag{3.1}$$

where FDI_{cit} is a measure of offshoring activity by U.S. multinationals, X_{cit} is a $k \times 1$ vector of other explanatory country-year, industry-year, and country-industry-year varying characteristics, α_0 and α_1 are scalar parameters, and β is a $1 \times k$ vector of parameters. The

[16] 2006 GSP Annual Review, written comment of Alcoa (USTR docket FR-0052)(2)

[17] Two examples from the recent 2011 Annual GSP Review include SC Johnson and Co., which petitioned for continued GSP benefits on behalf of its Thai subsidiary, and MAT Holdings, which petitioned on behalf of its production facilities in India. MAT Holdings spells out the mechanism plainly, writing "entering the United States duty-free under GSP also contributes to the success of U.S. investments abroad..." (Dockets USTR 2011-0015-0125 and -0083; USTR-2011-0015-0012)

[18] In principle, θ^*_{cjt} is not bound by the unit interval: negative values imply the U.S. would like to impose a tariff above MFN level, while values above 1 imply the U.S. would prefer an import subsidy.

[19] Due to partial-year eligibility, θ_{cjt} may take intermediate values when defined over a calendar year.

[20] Although the MNE data are available at the affiliate level, the financial and operating data are coded only to the NAICS-4 level.

parameters γ_c, δ_i, and η_t represent country-, industry-, and year-fixed effects, respectively. The remaining error term, ϵ_{cit}, represents unobserved heterogeneity in each country-industry-year observation and is assumed to be independent of both X_{cit} and FDI_{cit}. To the extent that the errors are correlated within countries or industries, country- and industry-fixed effects will mitigate the Moulton problem; to be conservative, we also cluster the standard errors by country in our benchmark specification.

We include in the vector X_{cit} the variables directly suggested by theory and a small set of standard political economy and gravity control variables. As discussed earlier, the key explanatory variables in X_{cit} include total U.S. domestic sales in industry i and year t and its square; total U.S. imports from the world in industry i and year t; total industry i exports to the U.S. from country c in year t; sales to the U.S. by affiliates of U.S. MNEs in the rest of the world (other than country c) to the U.S. in industry i and year t; and the U.S. MFN ad-valorem tariff rate in industry i and year t. We also include a set of industry-year controls to proxy for U.S. domestic political pressure, including the U.S. payroll, number of establishments, import penetration, number of employees, and the year-on-year log change in U.S. employment and import penetration. Two gravity variables, GDP per capita and population, vary at the country-year level.[21]

With a normally distributed error, the correct specification is a double-censored Tobit model. In practice, however, the three dimensions of fixed effects specified in (3.1) introduce a tradeoff between computational feasibility and adherence to this ex-ante preferred non-linear (Tobit) model specification. Thus, we pursue two different strategies for implementing our empirical test. In the first, we estimate a linear probability model in which we can remove industry, country, and year fixed effects, but must ignore the censoring process that generates the mass points for θ_{cit} at 0 and 1.[22] In a second approach, we adopt the non-linear double censored Tobit model with fixed effects entering as dummy variables, but the need to achieve convergence limits both our choice of estimator and the set of control variables we can include.[23]

[21] Our results are robust to the exclusion of these additional variables.

[22] Year-fixed effects are removed by dummy variables, while industry and country fixed effects are removed by demeaning. We correct the standard errors to account for the demeaning, but given the large number of observations, the correction is negligible.

[23] In the instrumented Tobit specification, we adopt the Newey (1987) two-step estimation and reduce the set of controls to only those variables explicitly indicated by theory.

We address the potential simultaneity between offshoring activity and preferential access by instrumenting for our measure of multinational affiliates sales of goods back to the United States. Guided by theory, we argue that market-seeking (horizontal) MNE sales to the local market should be independent of U.S. trade preferences.[24] At the same time, we expect local and offshoring MNE sales to be positively correlated because both reflect the attractiveness of the local market for foreign investors. Thus, we instrument for (export-oriented) sales of goods to the U.S. with (market-seeking) MNE affiliate sales (of goods and services) to the local market in the same country, industry, and year. As second instrument, we use worldwide sales of *services* by affiliates of U.S. MNEs, again within the same country, industry, and year. The same rationale applies to the second instrument: while service sales may reflect the attractiveness of the FDI host country within a given industry, we would not expect a direct link between affiliates' sales of services and U.S. preferences for goods trade. Crucially, we are careful to construct both instruments using *only those affiliates that do not also sell goods to the U.S.*, which addresses the potential concern that increasing returns to scale at the affiliate level could cause a firm's joint decision to enter and sell both goods to the U.S. and to the local market (or services worldwide) to depend directly on U.S. preferences.[25] Figure 1 summarizes our instrumentation strategy using a simple diagram.

Finally, to reduce the potential influence of outliers, we estimate all specifications taking logs of all variables that can take large integer values.[26] In baseline specifications, we also lag the explanatory right-hand side variables by one year to account for the time lag inherent to policy changes.

[24] One might be concerned that MNE sales to the local market could be a substitute for MNE sales to the U.S., and thus be correlated with the second-stage error term. The first-stage results suggest a strong positive correlation between sales to each destination, which argues against the substitutes story. The high p-values for the Hansen's J test offer additional reassurance.

[25] In our sample, 84% of affiliates report positive sales to the local market, while fewer than 21% sell goods to the U.S. Among affiliates that sell locally, fewer than 25% also sell goods to the U.S. Our results are robust to alternative instruments, including MNE affiliate sales to the ROW (rather than the local market), although we find this alternative instrument weaker and less plausibly exogenous given the greater potential substitutability across destinations for export-platform FDI. (Among the 35% of affiliates in our sample with positive sales to the rest of the world, 53% also sell goods to the U.S.)

[26] To preserve zeroes, we add one to (all) variables' values before taking the natural log (thus we replace x with $ln(1+x)$); the usual caveats apply. See, e.g., Wooldridge (2009), page 192.

4. Data

Preferential Market Access. The dependent variable of interest is the rate of preferential market access to the U.S. by industry, country, and year (θ_{cit}). One way to measure preferential market access is to use a direct measure of preference program eligibility. Using data from the U.S. Trade Representative (USTR), we code every HTS-8 tariff line by country and year as either eligible or ineligible for duty-free market access under an applicable U.S. preference program (NAFTA, GSP, AGOA, etc.). We then use time invariant trade weights (based on 1997 trade flows, just before our sample begins) to aggregate to the NAICS-4 industry level. The result is a weighted eligibility measure of duty-free access under any preference program, *Any Pref El*.[27]

While this simple eligibility-based definition of the preferential treatment variable is appealing in its parsimony, one can easily challenge the definition on the grounds that even when program eligibility is indicated by official trade policy, preferential treatment is often afforded to only a subset of the imports in question. Partial-year program eligibility and "competitive need limitations" (CNLs) are one concern, while restrictive rules of origin and other bureaucratic costs under some programs often make *de jure* preference eligibility useless in practice.[28]

With these caveats in mind, we define a second form of the dependent variable *Any Pref Share* using more detailed data from the U.S. International Trade Commission (USITC). Each year, the USITC reports the proportion of bilateral trade that clears U.S. Customs under each preferential program code, by industry and country of origin. We use this information to construct our baseline measure of the dependent variable, *Any Pref Share*, so that θ_{cit} is the (exact) share of country c exports of industry i in year t that entered U.S. Customs claiming duty-free access under any preferential program code. Note that this version of the dependent variable offers the additional advantage that it does not require an ad-hoc weighting scheme to aggregate to the NAICS-4 level.

[27] Our results are robust to removing the trade weights.

[28] Foreign exporters often do not claim preferential access when 'officially' eligible; see Hakobyan (2015) on underutilization of GSP preferences. The relevant concern for our study would be if MNE foreign affiliates are more likely to use preferential trade programs than non-affiliates (which is not obvious – disproportionate use of foreign-sourced inputs may make rules of origin harder to satisfy for MNE affiliates). The robustness of our results to either preference variable definition (usage or eligibility) suggests that this discrepancy is of minimal consequence in the context of our study.

In addition to these two measures of U.S. preferential market access under any preference program, we create versions of the dependent variables which only use GSP eligibility (*GSP El*) and usage (*GSP Share*). We find GSP preferences of particular interest not least because roughly two-thirds of country-industry-year observations with preferential market access are afforded duty-free treatment through GSP.[29] Moreover, the mechanics of the GSP program feature important institutional differences compared to regional or other preferential programs: firms and other interest groups regularly take part in the formal annual GSP review process, and there is considerable scope for discretion and annual variation in eligibility due to human rights, labor, intellectual property, or competitive need limitations violations. From an econometric perspective, the GSP-based preference measures are largely immune from concerns over reciprocal trade policy or bilateral investment provisions that could arise under other preference programs (i.e. Article XXIV free trade deals), and so may offer cleaner identification of a causal mechanism.

Finally, to convince the reader that there is sufficient variation left to explain after including country-, industry-, and year-fixed effects, we offer a set of simple plots in Figure 2.[30] On the vertical axis, we plot the residual of our preferential market access variables (*Any Pref Share*, *GSP Share*, *Any Pref El*, and *GSP El*) after controlling for country-, industry-, and year-fixed effects. On the horizontal axis, we plot country GDP per capita. In each plot, we see clearly that there is considerable variation in preferential market access under either the share or eligibility definitions, and that the degree of variation is higher among the less developed trading partners.

Multinational Affiliate Activity. The U.S. Bureau of Economic Analysis (BEA) collects detailed firm level financial and operating data for all foreign affiliates of U.S. multinational firms in which a U.S. entity holds an ownership interest of 10% or more.[31] We

[29] Weighting by trade volume, U.S. imports under GSP comprise only 7% of all preferential imports.

[30] A literal reading of GATT rules might lead one to expect that country and industry fixed effects should absorb most of the variation in trade preferences. (Article XXIV stipulates that among FTA partners, "virtually all" products should be eligible for duty-free access, while the Enabling Clause specifies that preferences be granted by industry and country, but not industry-country pair.)

[31] These data are collected in the BEA's benchmark and annual surveys of U.S. direct investment abroad for the purpose of producing aggregate statistics on U.S. multinational company operations for release to the general public. The confidential microdata that BEA maintains in its databases for research purposes are a by-product of its legal mandate to produce for the public aggregate statistics on multinational corporation operations.

necessarily restrict our sample to majority owned foreign affiliates (MOFAs), as only MOFAs report sales disaggregated by destination.[32] Affiliates are coded to NAICS-4 industries,[33] which we aggregate by country and year to create the variable FDI_{cit}: total sales of goods to the U.S. by MNE affiliates in country c, industry i, and year t, prorated by U.S. ownership.[34] For our instruments, we use the two additional measures from the BEA data discussed earlier: total sales to the local market by *only those MNE affiliates that do not also sell goods to the U.S.*, and total affiliate sales of services (worldwide), again by only those MNE affiliates that do not also sell goods to the U.S.

Control Variables. We include control variables at the country, industry, country-year, industry-year, and country-industry-year levels as discussed in the previous section. Sources and summary statistics are reported in Table 1.

Political Economy Measures. In several extensions, we explore the interaction between our key measures of domestic and multinational affiliate activity using explicit measures of the political organization of U.S. industries. Gawande and Bandyopadhyay (2000) develop an indicator variable, I^o, that takes a value of 1 for industries that are deemed politically active based on political contributions and import penetration ratios.[35] Bombardini (2008) develops an alternative proxy for political participation based on the within-industry dispersion of firm sales. We use two incarnations of her measure, ψ_{true}, which uses direct evidence of lobbying participation, and ψ_{sales}, which is based on sales dispersion alone.[36] Using the BEA data on MNE sales, we also construct ψ_{mne}, a direct analog to the (domestic) sales

[32] MOFAs constitute the majority (70% in 2007) of all foreign affiliate sales of U.S. MNEs, and virtually all sales when pro-rated by percentage of U.S. ownership.

[33] The BEA uses modified 'BEA NAICS' codes for industry categorization. Most codes are identical to the standard NAICS, but several are aggregations of standard NAICS (in which case we aggregate to concord the standard NAICS to the coarser BEA NAICS), and a few others are a disaggregation of the standard NAICS, in which case we concord the finer BEA NAICS to NAICS-4.

[34] BEA survey form BE-10B requires MNE affiliates to report sales disaggregated by goods vs. services for each destination market (local, U.S., or other). We map sales to NAICS industries according to each MNE affiliate's primary NAICS affiliation.

[35] Gawande and Bandyopadhyay (2000) regress PAC contributions on import penetration ratios interacted with 2-digit SIC dummy variables. If the coefficient on the interaction term is positive for a given SIC-2, it is assumed that all sub-industries within the SIC-2 category are politically active.

[36] Mapping to Bombardini's notation, $\psi_{sales} \equiv \theta(\rho = .5)$ while ψ_{true} is her "true theta" variable. See Bombardini (2008) for a detailed description of the construction of these measures.

dispersion measure, ψ_{sales}.[37] Descriptive statistics for these political economy measures are included at the bottom of Table 1.

Our data set spans 184 countries, 80 industries, and 10 years (1997-2006), which falls to a 9 year window when we introduce lags in the explanatory variables. In principle, we would have 132,480 observations, but because our dependent variable is defined as the preferential *share* of imports, we lose all country-industry-year observations for which U.S. imports are zero, leaving us with 68,130 observations in our benchmark sample.[38] When we consider the effect of GSP preferences specifically, we use a reduced sample that includes only those countries that are *de jure* GSP eligible. Table 1 summarizes the variables that we include in our full ($N = 68,130$) and reduced ($N = 42,849$) samples and lists their sources. A list of included countries and industries is available in the online appendix.

5. Results

Table 2 presents our benchmark estimation results. The first three columns report estimates from panel IV specifications, while columns 4-6 report results from non-linear IV Tobit specifications. We consider two definitions of the dependent variable: *Any Pref Share*, which is defined as the share of imports by industry i from country c in year t for which duty-free treatment was claimed under any trade preference program; and *GSP Share*, which is the share of imports by industry i from country c in year t for which duty-free treatment was claimed specifically under the U.S. GSP program. For the *Any Pref Share* specifications, we use two samples – the full sample, which includes all countries in our data set, and a reduced sample, which restricts attention to only (developing) countries that are *de jure* GSP eligible. For the *GSP Share* specifications, we use only the reduced sample.

In the panel IV specifications, country- and industry-specific fixed effects are removed by demeaning (with standard errors corrected accordingly); time-fixed effects are incorporated as year dummies; and standard errors are clustered at the country level. In the interest of space, the table reports coefficient estimates for only the variables of primary interest.[39]

[37] $\psi^{it}_{mne} \equiv .5\mu^{it}_{mne} + .5\sigma^{it}_{mne}$ where μ^{it}_{mne} and σ^{it}_{mne} represent the normalized mean and standard deviation of affiliate sales (worldwide) within industry i and year t.

[38] For consistency across specifications, we use this smaller data set in robustness checks of the eligibility-based dependent variable; our results are robust to using the larger data set.

[39] Explanatory variables included in the panel IV specifications but not reported are U.S. import penetration (ln), the annual log change of U.S. industry import penetration, number of U.S. establishments (ln), country population

The first column reports our benchmark case, estimating the effect of MNE goods sales to the U.S. on the rate of preferential access under *all preference programs granted to any country* in the full sample. Jumping immediately to the key result in the first row, we find that a 10% increase in multinational sales to the U.S. implies roughly a 4 percentage point increase in the share of imports receiving duty-free market access, which translates to a 20% increase in the rate of preferential access relative to the sample mean.[40] The result is statistically significant at the 5% level with clustering at the country level.

Reading down the table, all estimated coefficients for the remaining empirical predictions take the expected sign and are statistically significant at least at the 5% level, with the exception of the country-industry-year exports to the U.S., which are statistically insignificant. Consistent with political economy motives, we find that as a domestic industry grows relatively larger (as measured by U.S. domestic sales), it is more protected (offering less generous trade preferences), and that this relationship is both economically large and concave.[41] We also find that preferences become less generous as MNE affiliate sales from the ROW increase, which is once again consistent with our initial empirical predictions.

The theory did not offer clear priors for the sign of the remaining coefficients, but we venture that by at least a rough approximation, all of the estimates make sense. Our estimated coefficients are positive and negative for MFN tariffs and total U.S. imports, respectively. Tying back to the discussion in Section 2, the first result suggests that higher MFN tariffs may indeed shift political attention towards the use of preferences, while the second is consistent with the concern that increasing preferences for one country may reduce overall tariff revenue from the rest of the world via trade diversion. Finally, as one might expect, healthier U.S. industries with rising payrolls and faster year-on-year employment growth tend to offer more generous preferential access (or put another way, senescent industries receive more protection); at the same time, increasing the number of domestic employees (holding payroll fixed) seems to attract more protection in the form of a lower trade preference share.

Notice that the excluded instruments perform well in terms of explaining the potentially endogenous variable *MNE sales of goods to U.S.* The first stage results are as expected: the

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(ln), and GDP per capita (ln). Of these, only the country GDP per capita and the number of U.S. establishments are significant at the 10% level.

[40]From Table 1, the average rate of preferential market access is 19.8% in the full sample.

[41]Our coefficient estimates imply that the derivative of *Any Pref Share* with respect to U.S. domestic sales is negative for all values of domestic sales in our sample.

first stage coefficient estimates for the instruments *MNE local sales* and *MNE service sales* are positive and significant, and the F-statistic for the excluded instruments is 17.30 (p-value $< .001$).[42] Finally, the J-statistic p-value of 0.869 exceeds 0.1 by a substantial margin, passing at least this simple test for instrument validity.[43]

In column 2 of Table 2, we restrict the sample to exclude countries that are *de jure* GSP ineligible under the 1974 Trade Act.[44] In the reduced sample, 91% of observations apply to *currently* GSP eligible countries, opposed to 57% in the full sample. The effect of removing the *de jure* GSP ineligible countries is striking: under the otherwise identical specification, the estimated effect of vertical FDI more than triples compared to the full sample result in column 1. Among developing (potentially GSP-eligible) countries, a 10% increase in MNE sales to the U.S. leads to a 14.7 percentage point increase in the share of U.S. imports claiming any preferences. This effect is statistically significant at the 5% level, clustering at the country level.

In column 3 of Table 2, we reestimate the model using *GSP Share* as the dependent variable. Among potentially GSP-eligible countries, a 10% increase in MNE sales to the U.S. leads to a 14.9 percentage point increase in the share of U.S. imports claiming GSP preferences, statistically significant at the 1% level. A priori, one might have anticipated that trade preferences for poorer countries would be less responsive to MNE activity, but clearly we find no support for such a view.

The remaining variable coefficient estimates in columns 2 and 3 are consistent with our results in the full sample except that the effect of MNE affiliate sales from the ROW is

[42] The estimation results are very similar to the just-identified version of the model (not reported here), offering comfort that the over-identified model has not introduced systematic bias. An additional test for bias in over-identified models, comparing 2SLS results against LIML, also revealed no meaningful changes in the results. See Angrist and Pischke (2009), p. 209.

[43] As an additional test of instrument performance and potential reverse causality, we run a three-stage least squares (3SLS) specification in which our measure of FDI (MNE sales to the U.S.) is determined by trade preferences and our existing instruments. The coefficient estimate for the effect of *MNE Sales to the U.S.* on *Any Pref Share* remains remarkably similar at $.412^{***}$. Checking the reverse causality, we find that *MNE sales to the U.S.* depend positively and significantly (at the 1% level) on *Any Pref Share* (with a coefficient estimate of $.022^{***}$); *MNE local sales* ($.076^{***}$); and *MNE service sales* ($.427^{***}$). Complete results of this estimation are available from the authors upon request.

[44] Countries are deemed ineligible if they are members of the European Union, have high income according to the World Bank, are communist, have terrorist ties, or are members of an arrangement aiming at withholding supplies of vital commodities. The last two criteria appear to be implemented with considerable discretion, however, so we restrict the sample to exclude countries under only the first three criteria.

muted among developing countries, as one might expect to the extent that the trade diversion induced by a developing country is less likely to impose a competitive disadvantage on affiliate imports from the ROW.[45]

Finally, note that the instruments are somewhat weaker in the restricted sample, presumably because there are fewer industries among the developing countries with both market-seeking (local sales) affiliates and offshoring MNE affiliates selling goods back to the U.S. At the same time, instrument validity appears of little concern given J-statistic p-values of approximately 100% in these GSP-program specific specifications of the model.[46]

In the right half of Table 2, we turn to an IV Tobit specification for both the full and reduced samples, with double-censoring to restrict $\theta_{cit} \in [0, 1]$. As we noted earlier, demeaning the data to remove country- and industry-fixed effects is a legitimate exercise only in a linear framework, so we include country-, industry-, and year-fixed effects through dummy variables. Due to computational limitations in the estimation of the IV Tobit model with these three dimensions of fixed effects, we revert to a just-identified specification, reduce the set of controls to include only those explicitly specified by the theory, and resort to the Newey (1987) two-step estimation procedure to achieve convergence.[47]

Columns 4 and 5 of Table 2 report results for the share of imports entering the U.S. duty-free under any preference program for the full and the reduced sample, while column 6 provides results for the share of imports entering the U.S. duty-free under GSP for the reduced data set. The results are qualitatively similar to the results of the IV panel estimation, in that we find a strong positive effect of MNE sales on U.S. preferences, but the coefficient estimates are uniformly higher. For the *Any Pref Share* variable in column 4, the estimated effect of MNE sales to the U.S. more than triples compared to the IV panel specification, which suggests that in the full sample, our baseline linear specification may understate the effect of offshoring by multinationals on U.S. trade preferences. Reducing the sample to include only the *de jure* GSP-eligible countries in column 5 leads to roughly

[45] Among the not reported variable coefficients, per capita GDP loses statistical significance in the reduced sample.

[46] In the 3SLS specification of this regression, we find that *MNE sales to the U.S.* continue to depend positively and significantly (at the 1% level) on our instruments, *MNE local sales* (coefficient estimate .096***) and *MNE service sales* (.273***), but the reverse causality running from trade preferences (*GSP Share*) to vertical FDI is now negative and weaker (coefficient estimate $-.014^{**}$). Again, complete 3SLS results are available upon request.

[47] Unfortunately, Newey two-step estimation precludes clustering and limits the set of available post estimation statistics.

the same estimated effect of FDI on preferences as in the full sample, and the results are markedly closer to our linear model findings in the left half of Table 2, suggesting that the extent of downward bias in the linear model is much less acute in the reduced sample. The coefficient estimate is even higher if the dependent variable is the GSP share, which indicates that a 10% increase in MNE sales back to the United States is associated with a 20 percentage point increase in the rate of preferential access for countries and industries that already have some degree of preferential access at the NAICS-4 level. Finally, note that most of the censoring is at zero (55% of the observations for the *Any Pref Share* variable in the full sample) rather than 1 (just under 5% of the sample for the same dependent variable), which suggests that the lower bound at zero is likely responsible for bias in the linear specification results.

Summarizing our results thus far, we draw three broad conclusions. First, the empirical results are qualitatively consistent with the theory, and our instruments perform strongly in the full sample and only somewhat less so in the reduced sample. The key finding is that U.S.-bound sales by foreign affiliates of U.S. MNEs seem to increase the rate at which exports from the MNE host country and industry will be afforded preferential market access. Under the benchmark linear panel IV specification, a 10% increase in MNE sales to the U.S. is associated with an increase in preferential tariff exemption of about 20% of the mean for the sample of all countries.

Second, the effect seems to be much stronger among developing countries and under the GSP program. When we restrict the sample to exclude *de jure* GSP ineligible countries, we find that the estimates for both all preferences and GSP preferences rise more than threefold relative to the full sample findings. Moreover, in the reduced sample, the estimate for the effect of MNE sales on GSP preferences is slightly *higher* than that for all preference programs in general.

Lastly, the IV Tobit specification suggests that the linear IV panel estimates could represent a lower bound on the effect of MNE activity on preferential market access. In the IV Tobit specification, the effect of a 10% increase in MNE sales back to the U.S. entails a 15.4 percentage point increase in the rate of preferential market access under any program in the full sample and a 20 percentage point increase in the rate of preferential market access under GSP in the reduced sample.

6. Extensions and Robustness Checks

We now consider several extensions to the benchmark specification to further our understanding of the underlying mechanism and to test the robustness of our baseline results.

Political Economy Motivations. Given the potential for political lobbying to influence trade policy, we introduce interaction terms for both U.S. domestic sales and our measure of MNE offshoring with an indicator variable for whether or not an industry is politically organized, in the spirit of the Goldberg and Maggi (1999) and Gawande and Bandyopadhyay (2000) tests of the 'Protection for Sale' model (Grossman and Helpman, 1994). As described in Section 4, we use two distinct measures of political organization: the widely-used Gawande and Bandyopadhyay ('GB') binary variable designation of industry organization and a continuous political activity measure based on firm size dispersion developed more recently by Bombardini (2008).

Table 3 reports the results. The first three columns use the binary GB indicator for political organization, while the last two columns also incorporate firm-size distributions based on Bombardini (2008). Note that the sample size falls somewhat because the GB measures are not available for several industries.

Column 1 tests the performance of the GB measure for *domestic* demand for protection within our context: thus, we interact only U.S. domestic sales with the GB indicator of political organization, but leave the MNE sales measures unchanged from the baseline specification. Consistent with the existing literature, we find that the interaction between U.S. domestic sales and political organization is strongly negative and significant – suggesting that import protection (i.e. *less* preferential market access) is afforded to larger domestic industries *only if* they are also politically active.[48] Notice that the estimated effect of *MNE goods sales to U.S.* rises nearly twofold relative to the benchmark case, which suggests that by omitting explicit measures of domestic political organization, our benchmark specification may understate the effect of MNE activity on U.S. trade preferences. The remaining coefficient estimates are consistent with our earlier findings.[49]

[48] The coefficient estimate for the un-interacted U.S. domestic sales term is no longer statistically different from zero.

[49] In the interest of space, we suppress results for several of the control variables; the set of included controls is unchanged, all coefficient estimates are in line with earlier findings for the benchmark case.

The remaining columns of Table 3 reestimate the model including political economy interactions for both U.S domestic sales and the key measure of MNE offshoring activity; our implicit assumption is that the GB and Bombardini measures of domestic political organization also reflect political organization by multinational firms within the same industry. All else equal, we expect the interaction between MNE sales to the U.S. and political organization to be positive if indeed U.S trade preferences are differentially more responsive to politically active multinationals. Scanning the point estimates in the first and second rows of Table 3 in columns 2 through 5, we find evidence broadly suggestive that this is true.

In column 2, the coefficient estimate for MNE sales to the U.S. remains significant and positive, which suggests that our mechanism operates even for unorganized industries. The coefficient estimate for the interaction term is roughly of the same magnitude, which suggests moreover that politically organized industries are offered more than double the rate of preferential access for a marginal increase in MNE sales. Both coefficient estimates are positive and significant at the 5% level or more with robust standard errors. The remaining estimates are quantitatively very similar to the results in column 1, and need not be repeated. Column 3 reports the same results with standard errors clustered at the country level; statistical significance falls to roughly the 10% level, as one might anticipate given the weaker first stage with multiple endogenous regressors.[50]

Columns 4 and 5 repeat the exercise using measures of political organization based on Bombardini (2008). In column 4, we interact both U.S. domestic sales and MNE sales with ψ_{true}, which is the share of U.S. industry sales by firms above a certain size threshold (at least as large as the smallest politically active firm based on political contribution data) interacted with the GB measure of political activity. In column 5, we interact U.S. domestic sales with Bombardini's ψ_{sales}, which is the normalized average of the mean and standard deviation of firm sales within each U.S. industry (again, interacted with the GB measure of political organization). Using the BEA's firm level data on MNE affiliate sales, we create an MNE-specific analog to Bombardini's measure, ψ_{mne}, to interact with our measure of MNE sales. Both specifications using Bombardini's political economy measures deliver higher estimates for MNE sales absent the political economy interaction, with somewhat

[50]In columns 2 and 3, we add an additional instrument that interacts local MNE sales with the political organization measure; the F statistic falls somewhat as one might anticipate with two endogenous regressors. The Hansen's J statistic p-value remains sufficiently high not to trigger concerns over instrument validity.

weaker evidence of a differential trade policy response to MNE sales in politically organized industries. (Neither specification can withstand clustering at the country level; perhaps this is less surprising given that the continuous interaction term is noisier than the binary GB measure.) Interestingly, we find that the Bombardini measures suggest that *domestic* protection seekers (proxied by U.S. domestic sales) gain little additional influence relative to unorganized industries, counter to the findings with the GB measure.

While we remain cautious not to overstate the power of these results, they certainly suggest that politics – and in particular political lobbying – may play an important role in amplifying the mechanism. Moreover, they offer some additional evidence on the direction of causality driving our results. If the concern is that preferential market access is driving MNE sales to the U.S. (and not the reverse), it is hard to explain why the effect would depend on the political organization of the domestic industry.[51]

Falsification Tests. In Table 4, we offer additional evidence that U.S. trade preferences are responding to MNEs' sales back to the U.S., rather than some other correlated determinants of FDI that depend on the general attractiveness of the host country for foreign investors. For ease of comparison, we re-report the results from our benchmark specification with the full sample in the first column.

In column 2, we replace MNE goods sales to the U.S. by *only* the MNE goods sales to the U.S. that go to *affiliated* buyers (e.g. MNE parent or 'sibling affiliates').[52] Our prior is that these sales will have a stronger effect on trade preferences because more generous trade preferences are then beneficial for U.S. multinationals for two reasons: the foreign affiliates themselves will become more profitable and, moreover, some of the cost savings from lower tariffs may be passed onto the U.S. parent buying those imports. Indeed, in column 2 we observe that the coefficient on the new explanatory variable is slightly higher. However, this increase is quite small and the level of significance remains at 5% as before. From this, we conclude that the positive effect of export-platform FDI on trade preferences does not critically hinge on the idea that affiliate-produced inputs become cheaper for U.S. firms, but rather on the prospect of foreign MNE affiliates' increased profitability.

[51]We are grateful to an anonymous referee for this insight.

[52]Buyer affiliation is a BEA designation.

In column 3, we replace the key explanatory variable, MNE goods sales to the U.S., with MNE goods sales to the rest of the world (ROW) by all MNEs that are not also simultaneously selling goods to the U.S. For good measure, we additionally restrict attention to MNE sales to only unaffiliated (arms-length) buyers to avoid potential contamination by supply chain relationships within an MNE's multi-country umbrella. The falsification test is based on the idea that MNE sales to anywhere *other than* the United States should in principle have no effect on U.S. trade preferences or even a small negative effect (via general equilibrium effects), holding all else equal. In the third row of column 3, however, we find the estimated effect to be positive, albeit smaller and noisier. That said, we are not particularly surprised by this finding, since MNE affiliate sales are positively correlated across destinations: our positive result in column 3 might simply reflect omitted variable bias (the fact that we have not controlled for MNE sales to the U.S.). Once we reestimate the model controlling for U.S.-bound MNE sales in column 4, we indeed find the theoretical predictions validated: the coefficient on MNE goods sales to the U.S. is positive (and increased by roughly half compared to the baseline specification), whereas the effect of MNE goods sales to unaffiliated buyers in the ROW on trade preferences is smaller in size and negative.[53]

In column 5, we check whether the MNE sales to the local market (again, by firms that are not also selling goods to the U.S.) have any impact on U.S. trade preferences; i.e., we include one of our previously excluded instruments in the main specification. As replacement for the excluded instrument, we introduce MNE sales to unaffiliated buyers in the ROW. The coefficient of the previously excluded instrument, MNE sales to the local market, is very small and statistically insignificant. This result is in keeping with the exclusion restriction and offers more evidence that simply increasing the attractiveness of the local FDI-host country for foreign investors without increasing sales to the U.S. is not enough to obtain more preferential market access.[54]

Alternative Specifications and Dependent Variables. As an additional robustness check, we report results from a set of alternative econometric specifications. First, we

[53] In column 4, we instrument for MNE goods sales to the U.S. but not MNE sales to the rest of the world, as the latter satisfies the exclusion restriction. Estimates are little changed (qualitatively or quantitatively) if we instrument for both MNE sales variables.

[54] This result argues against the potential concern that U.S. trade preferences might be a reciprocal "favor" offered to FDI host countries in exchange for more favorable investment conditions.

consider two reduced form (no IV) linear versions of the model – one panel and the other pooled.[55] In a second robustness exercise, we consider the alternative definitions of trade preferences based on official program eligibility, which we aggregate to 4-digit NAICS using historical (1997) trade volumes.[56]

Table 5 reports the results using the full sample for the dependent variables *Any Pref Share* and *Any Pref El* and the reduced sample for *GSP Share* and *GSP El*. For brevity, the table lists only the coefficient estimate of interest – MNE sales to the U.S. – for each specification. Each cell in the table reports the key coefficient of interest for the model defined by the row label and the dependent variable definition designated by the column heading. For ease of reference, we also include the already reported baseline results from the linear panel IV and IV Tobit specifications.

Scanning the results in Table 5, the estimated relationship between MNE sales to the U.S. and trade preferences is broadly consistent across specifications and alternative definitions of the dependent variable. To evaluate the potential impact of alternative dependent variable definitions, we compare results across columns (1) and (2) for the effect of MNE sales on preferences granted under any program to any country, and columns (3) and (4) for the effect of MNE sales on GSP preferences offered to developing countries. Making these comparisons, we see that the estimates are generally similar across the two dependent variable definitions. The point estimates for the eligibility based measures of trade preferences are smaller and noisier than those for the share based variable definitions in the full sample, while the eligibility and share based variable definitions line up closely in the reduced sample. Finally, the *GSP* estimates are typically larger and more precisely estimated than the *Any Pref* estimates throughout.

Comparing the model specifications across rows, we note the panel instrumental variables results to be quantitatively consistent across specifications in both the full sample and even more so in the reduced sample. The OLS specifications yield smaller estimates, but remain positive and statistically significant. To summarize, while the coefficient estimates vary,

[55] We report significance for unclustered heteroskedasticity-robust and country-cluster robust standard errors in the panel OLS version and unclustered and two-way clustered (industry and country) standard errors in the pooled OLS version. The pooled version includes year fixed effects and time invariant controls: country distance and indicators for the textile and agricultural sectors.

[56] We also test unweighted aggregations of the preference eligibility variables and find qualitatively consistent results.

they are consistently positive and significant in most of the variants of the model reported in Table 5.

Pairwise Fixed Effects. Pushing the data in a different direction, we explore within-group variation more deeply by including pairwise fixed effects in Table 6. In columns 1 and 2, we rerun the baseline panel IV model for the full sample two more times: first with industry-year pair-fixed effects, country dummies, and the (instrumented) explanatory variable, MNE goods sales to the U.S., and then again with country-year pair-fixed effects, industry dummies, and (instrumented) MNE goods sales to the U.S. In each instance, we still find that the coefficient of our key explanatory variable of interest, MNE goods sales to the U.S., is positive and highly statistically significant, while the instruments continue to perform exceptionally well (with a strong first stage and high J statistic p-value).

The results for the reduced sample in columns 3 and 4 show once again that the impact of MNE goods sales to the U.S. is even more pronounced (higher by a factor of about 4 compared to columns 1 and 2) when we use the reduced sample and GSP share as the dependent variable. As in the IV panel specification, we observe that the instruments appear weaker, whereas the J statistic p-value increases.

7. Closing Remarks

In this paper, we examine the relationship between U.S. multinational affiliates and the structure of preferential tariff access to the United States. Combining firm level panel data on U.S. foreign affiliate activity from the U.S. Bureau of Economic Analysis (BEA) with detailed measures of implemented U.S. trade preferences from the U.S. International Trade Commission (USITC), we obtain a rich panel data set spanning 80 industries, 184 countries, and ten years (1997-2006).

Using instrumental variables to account for the endogeneity of export-oriented foreign investment, we show that more export platform FDI causes more generous trade preferences for goods originating from the country and industry towards which FDI is directed: within a given (4-digit NAICS) industry in a given country and year, a 10% increase in U.S. foreign affiliate exports to the U.S. is associated with roughly a 4 percentage point increase in the rate of preferential duty-free access from all preferential programs combined. We find moreover that GSP preferences, which are in principle designed to offer developing countries 'aid through trade' market access, seem to be particularly sensitive to MNE activities. Among countries eligible for the Generalized System of Preferences (GSP), a 10% increase in MNE sales to the U.S. is associated with a 14.9 percentage point increase in the share of goods claiming GSP preferences. We find this positive and significant relationship between U.S.-bound MNE sales and preferential treatment to be remarkably robust across a variety of model specifications and robustness checks.

As with any empirical study, there are caveats. The first is that our key explanatory variable, U.S.-bound MNE goods sales, almost certainly understates the extent of U.S. interests in the foreign export sector. Our data do not capture foreign arms-length suppliers to U.S. firms (multinational or otherwise) that may have as much sway in the preference setting process as the formal affiliates we measure in the BEA data. To the extent that we undercount the extent of U.S. MNE interests relative to the remaining variables, our quantitative estimates for the effect of MNE sales on U.S. trade preferences will be biased upward. By the same token, our results underestimate the political economy motive (if not the implementation) to the extent that WTO non-discrimination guidelines under Article XXIV and the Enabling Clause are binding. Absent these GATT limitations, our results suggest that discretionary application of trade preferences to favor a country's MNE affiliates would be even greater.

Even amid these qualifications, we conclude that discretionary trade preferences are influenced in part by the importing country's multinational firms' foreign direct investment decisions. The results are strong; the implication is clear: investment matters. Whether this policy sensitivity is good or bad depends on how the interplay between FDI, multinational firms, and trade policy manifests in practice. For some trading partners, generous preferential tariff treatment may foster additional export-oriented investment, reinforcing a virtuous cycle of improved market access and increased FDI. To the extent that the same mechanism induces significant investment and trade diversion, however, other countries may be left out entirely. How these competing effects play out in coming years will remain an important question for researchers and policy makers alike.

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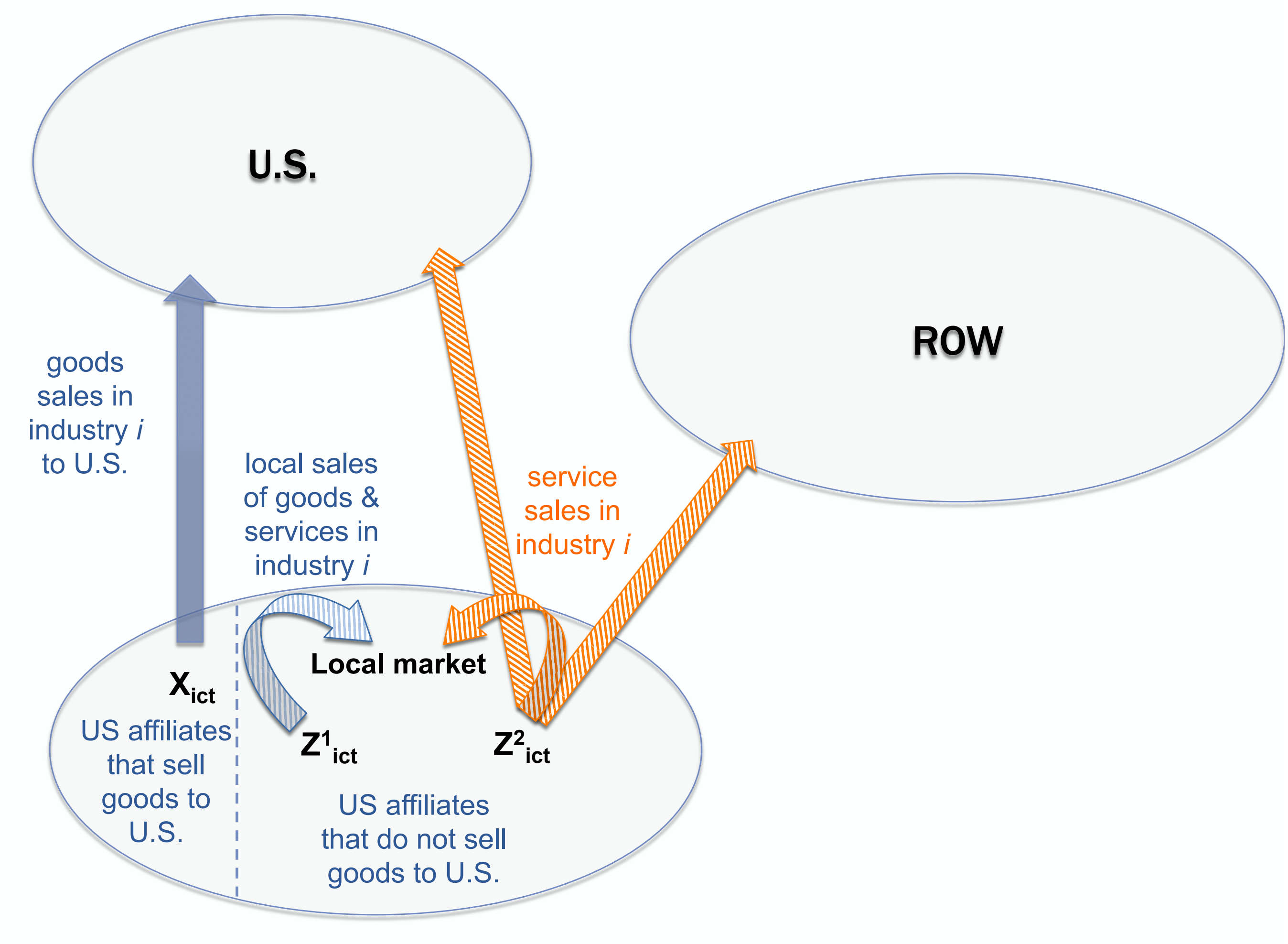

FIGURE 1. Sketch of Instrumentation Strategy

We construct two instruments for sales of goods to the U.S. by affiliates of U.S. MNEs, denoted X_{ict} above.

- Instrument Z^1_{ict} represents sales to the local market (of goods and services) by only those affiliates of U.S. MNEs that do not also sell goods to the U.S. market.
- Instrument Z^2_{ict} includes worldwide sales of services (only), again by only those affiliates of U.S. MNEs that do not also sell goods to the U.S.

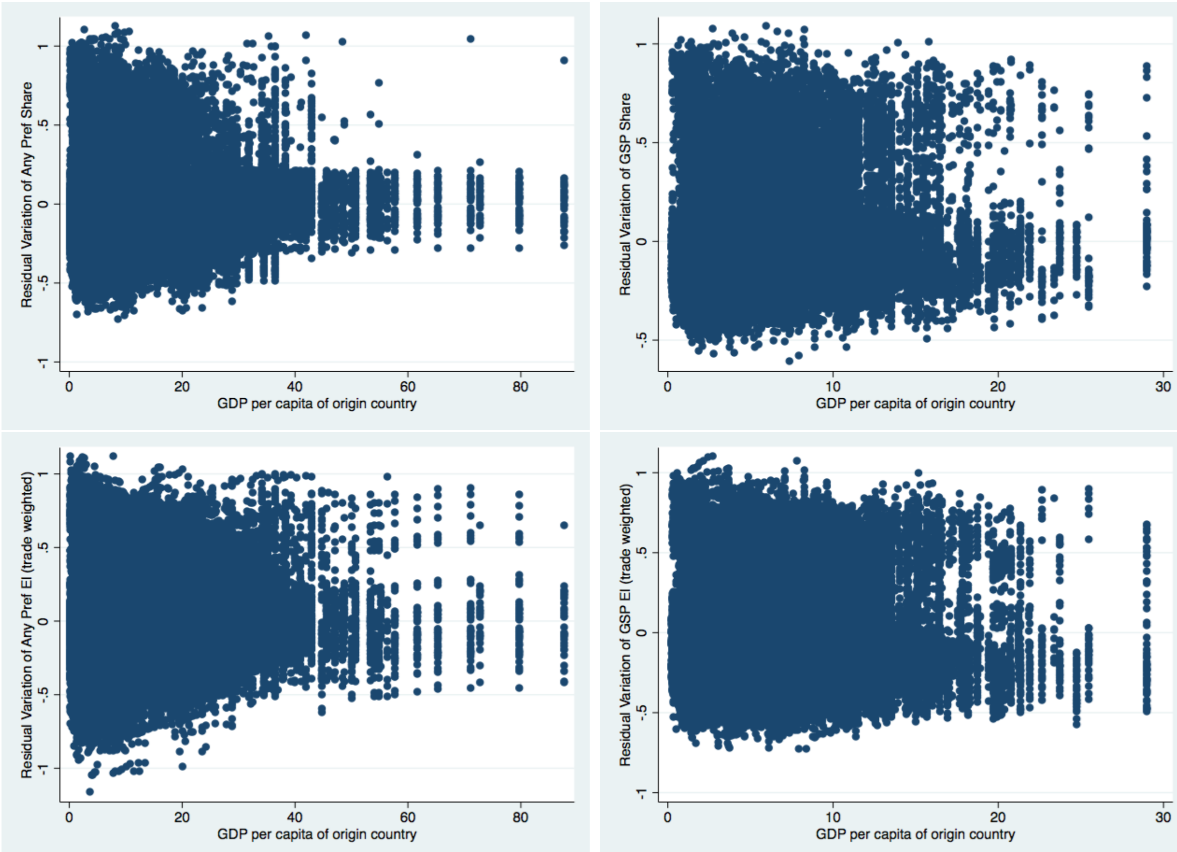

FIGURE 2. Residual Variation in Trade Preferences after Removing Country-, Industry-, and Year-Fixed Effects

Table 1. Data Summary Statistics

	Full Sample				Reduced Sample	
Variable; Source	Mean	SD	Min	Max	Mean	SD
distance to U.S. (km); www.timeanddate.com	8,305	3,681	734	16,357	8,297	3,698
terrorist; CIA World Fact Book	.013	.115	0	1	.021	.144
communist; CIA World Fact Book	.021	.143	0	1	0	0
population (b.); Penn World Tables (PWT)	.051	.168	.00002	1.314	.046	.136
per capita GDP (USD); PWT	13,547	11,776	288.4	87,825	6,531	4,604
textile indicator	.058	.234	0	1	.066	.249
agriculture indicator	.035	.184	0	1	.041	.197
U.S. employees (m.); US Census	.318	.550	.015	3.47	.343	.585
U.S. payroll (mUSD); US Census	8.69	7.06	.678	35.3	8.99	7.15
U.S. sales (mUSD); US Census	62.3	54.1	5.03	546.8	63.5	52.8
U.S. total imports (bUSD); USITC, USTR	17.2	23.2	.090	214.7	17.8	22.8
MFN ad-valorem eqv. (wt); USITC, USTR	.027	.039	0	1.36	.029	.043
U.S. num. establishments (th.); US Census	41.9	192.1	.092	1,151	47.7	205.3
U.S. import penetration	.990	.017	.600	1.00	.990	.015
log change U.S. employees	−.031	.061	−.494	.145	−.032	.061
log change U.S. import pen.	.001	.005	−.020	.176	.001	.004
c- i- t- exports to U.S. (bUSD); UN Comtrade	.161	1.18	$2.5 \text{x} 10^{-7}$	59.2	.047	.400
country curr. GSP el.; USTR	.573	.495	0	1	.910	.287
country *de jure* GSP inel.; USTR	.371	.483	0	1	0	0
industry curr. GSP el.; USITC, USTR	.996	.064	0	1	.997	.057
Any Pref Share; USITC, USTR	.198	.340	0	1	.276	.377
GSP Share; USITC, USTR	.122	.275	0	1	.193	.326
El Any Pref (hwt); USITC, USTR	.288	.392	0	1	.348	.422
El GSP (hwt); USITC, USTR	.206	.367	0	1	.327	.416
MNE goods sales to U.S. (bUSD); BEA	.027	.515	0	*D*	.0063	.1271
MNE goods sales to U.S. (prorated by ownership)	.026	.509	0	*D*	.0062	.1268
(instr.) MNE local sales† (bUSD); BEA	.041	.245	0	*D*	.014	.111
(instr.) MNE services sales† (mUSD); BEA	.349	8.67	0	*D*	.087	3.31
any MNE; BEA	.307	.461	0	1	.198	.399
rest of world MNE goods sales to U.S. (bUSD); BEA	2.43	6.11	0	*D*	2.41	5.72
political org., I^o; Gawande-Bandyopadhyay (2000)	.472	.499	0	1	.466	.499
political dispersion, ψ_{true}; Bombardini (2008)	.270	.173	0	.866	.264	.171
political dispersion, ψ_{sales}; Bombardini (2008)	.041	.064	.005	.634	.038	.056
political dispersion, ψ_{mne}; BEA	.017	.060	0	1	.011	.049

68130 (42849) observations in full (reduced) sample; ‘D’: BEA data redacted for confidentiality, used in analysis.

† Indicates sales by *only* those MNE affiliates that are not also selling goods to the U.S.

Table 2. Panel IV and IV Tobit Results

	Panel IV Results			IV Tobit Results		
	Any Pref Share		GSP Share	Any Pref Share		GSP Share
	(1)	(2)	(3)	(4)	(5)	(6)
MNE goods sales to U.S.	.409**	1.47**	1.49***	1.54***	1.57***	2.01***
[ln, billions USD]	(.204)	(.620)	(.575)	(.285)	(.547)	(.541)
U.S. domestic sales	-.897***	-1.34***	-.971***	−.297	−.774	2.04***
[ln, billions USD]	(.238)	(.400)	(.316)	(.557)	(.705)	(.723)
U.S. domestic sales, sq.	2.94***	4.74***	3.04***	4.04***	6.56***	−.953
[ln, billions USD]	(.479)	(.843)	(.661)	(1.23)	(1.55)	(1.72)
c- i- t- exports to U.S.	-.066	-.346***	-.401***	−.163**	−.215*	−.405***
[ln, billions USD]	(.046)	(.098)	(.096)	(.064)	(.115)	(.113)
ROW MNE sales to U.S.	-.009**	-.007	-.003	.004	−.004	−.006
[ln, billions USD]	(.004)	(.006)	(.005)	(.009)	(.012)	(.012)
U.S. total imp. (all countries)	-.036***	-.052***	-.059***	−.124***	−.157***	−.189***
[ln, billions USD]	(.006)	(.010)	(.008)	(.023)	(.030)	(.029)
MFN ad-valorem tariff	1.62***	2.03***	1.11***	2.92***	3.33***	2.24***
[ln]	(.181)	(.204)	(.162)	(.097)	(.115)	(.110)
U.S. employees	-.175***	-.267***	-.176***			
[ln, millions]	(.043)	(.059)	(.047)			
U.S. payroll	.031**	.041**	.084***			
[ln, millions USD]	(.014)	(.020)	(.015)			
log change U.S. employment	.099***	.118***	.063***			
[year-on-year]	(.019)	(.028)	(.024)			
additional controls‡	yes	yes	yes	no	no	no
country, year, & industry FEs	yes	yes	yes	yes	yes	yes
instrument: MNE local sales [ln]†	yes	yes	yes	yes	yes	yes
instrument: MNE service sales [ln]†	yes	yes	yes	no	no	no
KP Wald F-stat	17.30	3.91	3.91			
AR Wald test χ^2 p-value	.124	.004	$< .001$			
Hansen's J stat p-value	.869	.875	~ 1.0			
cluster variable	country	country	country			
left censored obs. (dep. var.= 0)				37,612	19,410	22,724
right censored obs. (dep. var.= 1)				3,122	2,997	1,953
sample	Full	Reduced		Full	Reduced	
observations	68,130	42,849		68,130	42,849	

Heteroskedasticity-robust standard errors in parentheses. ***, **, * denote significance at 1%, 5%, and 10% levels, resp. ‡ Additional controls: (industry-year) U.S. import penetration (ln), annual log change of U.S. industry import penetration, number of U.S. establishments (ln), (country-year) GDP per capita (ln), population (ln).

† Instruments include MNE sales by *only* those MNE affiliates that are not also selling goods to the U.S.

TABLE 3. Political Economy Interactions

	Gawande-Bandyopadhyay			Bombardini	
Dependent variable: Any Pref Share	(1)	(2)	(3)	(4)	(5)
MNE goods sales to U.S.	.731**	.413**	.413*	.590***	.669***
[ln, billions USD]	(.313)	(.173)	(.230)	(.175)	(.184)
MNE sales to U.S. × Pol. Org		.598***	.598	.470*	.921
[ln, billions USD]		(.159)	(.376)	(.240)	(.822)
U.S. domestic sales	-.281	-.127	-.127	-.928**	-.995**
[ln, billions USD]	(.364)	(.388)	(.363)	(.400)	(.400)
U.S. domestic sales × Pol. Org.	-1.26***	-1.44***	-1.44***	-.317	-1.52
[ln, billions USD]	(.272)	(.279)	(.280)	(.530)	(3.31)
U.S. domestic sales, sq.	1.69**	1.51**	1.51**	2.94***	3.04***
[ln, billions USD]	(.659)	(.766)	(.647)	(.789)	(.791)
c- i- t- exports to U.S.	-.110**	-.116***	-.116**	-.108***	-.112***
[ln, billions USD]	(.052)	(.038)	(.052)	(.030)	(.030)
ROW MNE sales to U.S.	-.007	-.004	-.004	-.005	-.005
[ln, billions USD]	(.006)	(.006)	(.006)	(.006)	(.006)
U.S. total imp. (all countries)	-.028***	-.029***	-.029***	-.028***	-.027***
[ln, billions USD]	(.009)	(.010)	(.009)	(.010)	(.010)
MFN ad-valorem tariff	1.66***	1.66***	1.66***	1.66***	1.66***
[ln]	(.190)	(.082)	(.191)	(.082)	(.082)
...					
Pol. Org. for U.S. domestic sales	I^o	I^o	I^o	ψ_{true}	ψ_{sales}
Pol. Org. for MNE sales	–	I^o	I^o	ψ_{true}	ψ_{mne}
additional controls‡; c-, i-, and t- FEs	yes	yes	yes	yes	yes
Instrument: MNE local sales† [ln]	yes	yes	yes	yes	yes
Instrument: MNE service sales† [ln]	yes	yes	yes	yes	yes
Instrument: MNE local sales† [ln] × Pol. Org.	no	yes	yes	yes	yes
KP Wald F-stat	12.34	11.28	5.542	16.20	16.49
AR Wald test χ^2 p-value	.065	$< .0001$	.027	$< .0001$	$< .0001$
Hansen's J stat p-value	.913	.375	.352	.567	.723
cluster variable	country	–	country	–	–
number of clusters	184	–	184	–	–
observations	60,704	60,704	60,704	60,704	60,704

Heteroskedasticity-robust standard errors in parentheses. ***, **, * denote significance at 1%, 5%, and 10% levels, resp. ‡ Additional controls: U.S. employees (ln), payroll (ln), import penetration (ln), annual log change of U.S. employment and U.S. industry import penetration, number of U.S. establishments (ln), per capita GDP (ln), population (ln).

† Instruments include MNE sales by *only* those MNE affiliates that are not also selling goods to the U.S.

Table 4. Mechanism and Falsification Tests

Dependent Variable: *Any Pref Share*	(1)	(2)	(3)	(4)	(5)
MNE goods sales to U.S.	.409**			.636**	
[ln, billions USD]	(.204)			(.293)	
MNE goods sales, only aff. buyers in U.S.		.479**			
[ln, billions USD]		(.237)			
MNE goods sales, only unaff. buyers in ROW†			.318*	-.181**	
[ln, billions USD]			(.176)	(.081)	
MNE sales to local market†					-.027
[ln, billions USD]					(.054)
U.S. domestic sales	-.897***	-.905***	-.874***	-.908***	-.826***
[ln, billions USD]	(.238)	(.239)	(.234)	(.244)	(.238)
U.S. domestic sales, sq.	2.94***	2.95***	2.95***	2.93***	2.86***
[ln, billions USD]	(.479)	(.482)	(.470)	(.492)	(.475)
c- i- t- exports to U.S	-.066	-.070	.019	-.114*	.027*
[ln, billions USD]	(.046)	(.047)	(.015)	(.059)	(.016)
U.S. total imp. (all countries)	-.036***	-.035***	-.039***	-.035***	-.038***
[ln, billions USD]	(.006)	(.006)	(.007)	(.007)	(.006)
ROW MNE sales to U.S.	-.009**	-.010**	-.010***	-.009*	-.009**
[ln, billions USD]	(.004)	(.004)	(.004)	(.005)	(.004)
MFN ad-valorem tariff	1.62***	1.61***	1.63***	1.61***	1.63***
[ln]	(.181)	(.180)	(.185)	(.180)	(.185)
...					
additional controls‡; c-, i-, and t- FEs	yes	yes	yes	yes	yes
Instrument: MNE local sales† [ln]	yes	yes	yes	yes	
Instrument: MNE service sales† [ln]	yes	yes	yes	yes	yes
Instrument: MNE unaff. sales ROW †[ln]					yes
KP Wald F-stat	17.30	13.02	20.72	10.68	59.22
AR Wald test χ^2 p-value	.124	.124	.124	.040	.092
Hansen's J stat p-value	.869	.978	.643	.976	.078
cluster variable	country	country	country	country	country
number of clusters	184	184	184	184	184
observations	68,130	68,130	68,130	68,130	68,130

Heteroskedasticity-robust standard errors in parentheses. ***, **, * denote significance at 1%, 5%, and 10% levels, resp. ‡ Additional controls: U.S. employees (ln), payroll (ln), import penetration (ln), annual log change of U.S. employment and U.S. industry import penetration, number of U.S. establishments (ln), per capita GDP (ln), population (ln).

† Instruments include MNE sales by *only* those MNE affiliates that are not also selling goods to the U.S.

Table 5. Alternative Specifications: Results Summary

Coefficient Estimates for MNE Goods Sales to U.S. (prorated by U.S. ownership)

	Full Sample		Reduced Sample	
	Any Pref Share	Any Pref El	GSP Share	GSP El
	(1)	(2)	(3)	(4)
Panel IV [unclustered/country clustering]	.41***/**	.29**/−	1.49***/***	1.61***/−
IV Tobit [with c-, i-, t- fixed effects]	1.54***	.31	2.01***	1.14*
Panel OLS [unclustered/country clustering]	.08***/***	.02*/−	.30***/***	.34***/***
Pooled OLS [unclustered/2-way clustering]	.14***/***	.09***/**	.22***/**	.29***/**

***, **, * denote significance at 1%, 5%, and 10% levels, respectively.

Table 6. Pairwise Fixed Effects

	Full Sample (Any Pref Share)		Reduced Sample (GSP Share)	
	(1)	(2)	(3)	(4)
First Stage				
MNE local sales†	.150***	.151***	.125***	.129***
[ln, billions USD]	(.014)	(.014)	(.031)	(.030)
MNE sales of services†	.645	.623	.293	.275
[ln, billions USD]	(.391)	(.387)	(.572)	(.572)
Second Stage				
MNE goods sales to U.S.	.230***	.234***	.983***	1.05***
[ln, billions USD]	(.053)	(.049)	(.283)	(.290)
observations	68,130	68,130	42,849	42,849
industry-year; country fixed effects	Y	N	Y	N
country-year; industry fixed effects	N	Y	N	Y
First-stage KP F-stat	68.82	70.11	9.16	9.85
First-stage AR Wald χ^2 p-value	$< .0001$	$< .0001$	$< .0001$	$< .0001$
Hansen's J stat p-value	.558	.869	.990	.988

Heteroskedasticity-robust standard errors in parentheses. ***, **, * denote significance at 1%, 5%, and 10% levels of significance, respectively. In each specification, pairwise fixed effects are removed by demeaning, while the second dimension of fixed effects is included via dummy variables. Instruments include MNE local sales and sales of services by *only* those MNE affiliates that are not also selling goods to the U.S.