U.S. MULTINATIONALS AND PREFERENTIAL MARKET ACCESS

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ABSTRACT. This paper examines the relationship between offshoring activity by U.S. multinational firms and the structure of U.S trade preferences. We combine firm level panel data on U.S. foreign affiliate activity from the U.S. Bureau of Economic Analysis (BEA) with detailed measures of U.S. trade preferences from the U.S. International Trade Commission (USITC) to create a three-way panel that spans 80 industries, 184 countries, and ten years (1997-2006). Consistent with existing theory, we find that offshoring multinational activity and preferential market access are positively and consistently correlated, both in the pooled sample and within countries, industries, and years. Using both instrumental variables and simultaneous equations approaches to address the likely endogeneity of export-oriented foreign investment, we find that each $1 billion in U.S. foreign affiliate exports to the U.S. from a particular industry and country is associated with roughly a 3.5 percentage point increase in the rate of preferential duty free access. Restricting attention to the Generalized System of Preferences (GSP), the dollar-for-dollar influence of multinational affiliate sales on preferential market access declines by roughly a third for the overall sample, but rises by more than an order of magnitude when we restrict attention to potentially GSP eligible countries.

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1. Overview

Recent theoretical work 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 its MNEs' 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 tailor their commercial policies in response to the interests of constituent industries (particularly in the presence of lobbying pressure), differences in the pattern of firm operations across the globe may be reflected in trade policy.

These ideas are formalized in Blanchard (2007) and (2010), which evaluate the implications of international investment for trade negotiations. Blanchard (2010) explores how cross-border ownership – in any sector and by any country – may translate into an altered role for multilateral tariff negotiation under the auspices of the GATT and its successor institution, the World Trade Organization (GATT/WTO). Blanchard (2007) uses a more specialized model to demonstrate how a large country's overseas investments can influence its government's optimal tariff policy towards the FDI-host countries when international investment is endogenous and therefore dependent on trade policy. The paper extends the model to a multi-country setting to argue that the possibility of preferential tariff arrangements can prove an effective means for harnessing the trade liberalizing potential of foreign direct investment if (but only if) international investment is of the export-oriented type.

Unfortunately, these simple empirical predictions are not easily taken to the data, which may explain why so little work has been done in this regard.[1] Empirically testing the hypothesis that cross-border investment influences governments' most preferred trade policies proves problematic first and foremost because most advanced economies

[1]The notable exception is recent work by Blonigen and Cole (2010) who link pre-WTO accession tariffs in China to FDI.

have set tariffs cooperatively since the inception of the General Agreement on Tariffs and Trade (GATT) in 1947. Since tariff concessions are negotiated multilaterally through coordinated rounds and are subject to the Most Favored Nation (MFN) non-discrimination clause, the econometrician is challenged to identify the influence of the pattern of foreign direct investment (FDI) apart from other, often unobservable yet likely predominant, multilateral pressures at the negotiating table. A government is unlikely to change its MFN tariff on a particular good (reducing the tariff for all of its MFN trading partners) if its underlying objective is to improve market access for its foreign affiliates in just a handful of countries. Similarly, 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, again diluting the potential for MNE offshoring activity in a particular industry and country to influence the MFN tariff.

Our empirical strategy sidesteps these difficulties by focusing on the potential influence of MNE activity not on MFN tariffs, but on the recent proliferation of various preferential trade agreements and the Generalized Systems of Preferences (GSP) by which industrial nations grant developing countries facilitated market access. Preferential treatment of trade flows is exempt from MFN (under Article XXIV of the GATT and the Enabling Clause for FTAs and GSP, respectively), and therefore may be considered a closer reflection of a government's unilateral trade preferences.[2]

The second potential complication for empirical testing lies in differentiating export-oriented (vertical) FDI apart from import-substituting (horizontal) FDI. While theory predicts that export-oriented FDI will exert downward pressure on tariffs in the investment-source country, the converse holds for import-competing investment. To the

[2]While in principle Article XXIV and GSP limit the extent of unilateral discretion, the intended uniformity (across industries and/or countries) of MFN exemptions appears to be quite weak in practice, as shown in Figure 1. See Blanchard and Hakobyan (2012) for a detailed analysis of the extent of discretion exercised under the U.S. GSP. Indeed, were FTAs truly non-discriminatory across industries – or were GSP preferences in practice determined only by country and industry (but without discretion at the level of the country-industry pair or year-to-year variation) – then our results would not withstand country- and industry-fixed effects, as they do.

extent that multinationals operate horizontal 'tariff jumping' operations abroad, those activities will have either a negligible or small *positive* effect on the investment-source country's optimal tariff.[3] Indeed, import-competing FDI and trade protection are often positively correlated in practice, largely due to the reverse causality as firms circumvent protectionist barriers *abroad* by building tariff jumping factories in their target market.[4] Fortunately, the richness of the BEA data offers us an empirical solution. In our data, MNE sales are disaggregated by destination; we are thus able to distinguish export-oriented investment (measured as U.S. foreign subsidiaries' goods sales to the U.S.) apart from horizontal import-substituting investment (subsidiaries' sales to the foreign local market).[5]

The last and biggest hurdle in identifying a potential effect of MNE activity on trade policy is the clear endogeneity of export platform investment. A potential FDI host country becomes a more attractive venue for offshoring operations if it has preferential market access to the anticipated market for its exports. Mexico's Maquiladora program, the well known North American Free Trade Agreement (NAFTA) predecessor, is an obvious example: duty free access to the U.S. market was precisely the carrot offered to entice investors to set up export-oriented manufacturing bases south of the border. Indeed, as the theory laid out formally in Blanchard (2007) makes quite clear, 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. While import-competing horizontal investment should not itself influence or be influenced by U.S. tariffs (as argued above), it is positively correlated with export-oriented investment (presumably because both rely on a favorable

[3]An increase in the investment source country's tariff will cause the world relative price of the foreign import good (and thus the return to horizontal FDI) to rise.

[4]Recall the prominent example of 'tariff jumping' Japanese car manufacturing plant construction in the U.S. in response to the voluntary export restraints in the 1980s.

[5]The BEA data include a third category: MNE sales to the rest of the world. Although these sales are clearly also export platform, they would not benefit directly from improved access to the U.S. market.

climate for investment). Thus, we use (import-competing) MNE sales to the local market as an instrument for (export-oriented) MNE sales to the U.S.

For this paper, we assemble a three-way panel data set including 80 industries, 184 countries, and ten years (1997-2006) to answer the question of whether export-oriented operations by U.S. MNEs cause higher rates of trade preferences for imports originating from countries where U.S. firms have set up shop. Our findings are consistent with the presence of such a causal relationship: conservatively, an additional $1 billion of U.S. foreign affiliate exports to the U.S. (roughly a one standard deviation increase) is associated with an increase in the duty free access rate of about 3.5 percentage points, controlling for the endogeneity of export-oriented FDI; this effect proves to be remarkably robust in a variety of different empirical specifications.

Our empirical results provide compelling evidence that offshoring MNE activity spurs preferential trade liberalization to the MNE's home country, which further deepens economic integration between the investment host and investment source countries. To the extent that more generous preferential tariff treatment fosters additional export-oriented investment, the cycle of improved market access and increased FDI may continue. At the same time, however, it stands to fear that the same mechanism can lead to substantial trade and investment diversion; a particular concern is that just as some trading partners experience ever-greater economic integration through this investment-trade nexus, other countries may be left out entirely.

The remainder of the paper proceeds in the usual sequence. In the next section, we briefly relate this paper to earlier work. Section 3 then sketches the basic theory, which Section 4 molds into our empirical strategy. A description of the data follows in Section 5, while Section 6 presents the results. We describe a series of robustness tests in Section 7 before concluding in Section 8.

2. Related Literature

Our study complements and considerably extends the empirical literature on the determinants of preferential treatment and the influence of international investment. The literature on trading blocs is sufficiently broad that discussion is restricted here to the small subset of work most closely related to this paper.[6] Most relevant to this paper's objective are Magee (2003) and Baier and Bergstrand (2004), which test empirically the determinants of preferential trade agreements. Both papers find qualitatively similar results: trade agreements tend to form between countries that are similar in size, geographically proximate, and politically liberal. The close concordance of their results is particularly striking given that these papers were written concurrently but independently from each other (each uses a distinct empirical strategy and data set). Both Magee (2003) and Baier and Bergstrand (2004) also have in common that they view preferential treatment as a binary all-or-nothing decision between two countries and do not exploit, as we do, the considerable variation in preferential treatment at the industry level over time.

Also related is DeVault (1996), which examines whether U.S. trade preferences during 1988-1994 for country-product pairs within the GSP framework are explained by domestic import-competing industry and exporting country characteristics. Along a similar vein, Lederman and Özden (2007) and Özden and Reinhardt (2005) recognize and explore the geo-political determinants of U.S. trade preferences in the process of examining the effect of trade preferences on developing countries' trade and preference beneficiaries' export patterns. Kee, Olarreaga, and Silva (2007) investigate whether foreign lobbying increases preferential market access to the U.S. in a Protection for Sale framework where the government values both foreign lobby contributions and domestic welfare, but with differing weight. Finally, in an important recent contribution, Ludema, Mayda, and Mishra (2011) examine the extent to which U.S. tariff suspensions (relatively small but highly discretionary exemptions to MFN tariffs) respond to both monetary and non-monetary lobbying by U.S. firms. Summarizing, while each of these earlier studies captures important elements of preferential tariff setting, none consider the potential influence of international investment.

[6]See Bhagwati, Krishna, and Panagariya (1999) for a broader review.

A final subset of related work consists of a handful of studies that examine the role of (exogenous) PTAs in determining investment flows. The seminal theoretical work by Motta and Norman (1996) identifies the potential for PTA member countries to attract both export-oriented and import-competing investment from investors within and outside a PTA, depending on the external barriers to the trading bloc. Balasubramanyam, Sapsford, and Griffiths (2002) offer a first test for the potential influence of regional integration on investment flows, treating preferential trade agreements ("regional integration areas" or "RIAs," in their lexicon) as exogenous. Using cross-section data for 1995 on the aggregate bilateral investment flows for 381 country pairs, they find that once standard gravity variables and country characteristics are included in the estimation, the presence of RIAs has no predictive power for FDI flows. 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 project provides an important complement to a companion paper, Blanchard and Matschke (2006). There, we use U.N. data on aggregate bilateral investment positions to predict the future formation of free trade agreements. Although our earlier study offers a wide geographic scope, spanning 37 countries and several decades, it suffers from systematic data deficiencies borne of the large coverage area and multiple reporting agencies. (The bilateral investment position data from each country are self-reported to the United Nations (UN) and exhibit substantial cross-country discrepancies due to different accounting methodologies. Moreover, since only 12 of the 178 reporting UN members have *ever* provided the UN with investment data disaggregated by industry, only aggregate measures of bilateral investment could be used in the analysis.) A key finding from that project is that we should focus on using industry-level investment data to better estimate the effects of vertical (export-oriented) investment apart from the influence of horizontal (import-sector) investment. By using detailed U.S. data on MNE sales in this study, which are provided at the industry level and separated by destination of sales (U.S., local, or rest of the world), we are able to identify the effects of investment in import-oriented versus export-oriented sectors of the economy. Moreover, the U.S. BEA-DIA data has the additional benefit of consistent reporting across countries and years.

3. Theoretical Framework

To present the underlying theory, we use a simplified version of Blanchard (2007). For the purpose of this paper, we assume that the country's aggregate utility is quasi-linear, which allows us to conduct a partial equilibrium analysis for any non-numeraire good.

Consider a country that charges a common ad-valorem MFN tariff on the imports of a given non-numeraire good from some finite number of foreign trading partners. The Home country is bound by its ad-valorem MFN tariff τ on the good, but can, for each trading partner c, choose the share of imports that is exempt from MFN, θ_c. The Home country's total welfare from income and consumption associated with the good can be written:

$$W = \alpha_d \pi(p) + V(p) + \sum_{c \in N} [(1-\theta_c)\tau p^W M_c(p, \theta_c) + \alpha_c r_c^*(\theta_c)\hat{K}_c], \tag{3.1}$$

where N is the set of foreign trading partners; α_i denotes the welfare weight on returns to Home-owned producers located domestically (α_d) or abroad in country c (α_c); π is domestic profit; V is consumer surplus; M_c is imports from country c; and $r_c^*\hat{K}_c$ denotes rental income from FDI in country c. The domestic price is denoted by p and the world market price of the good by p^W. Domestic profits and FDI rental returns from overseas investments may receive welfare weights greater than 1 due to political economy influences (i.e. if $\alpha_d, \alpha_c \geq 1$). We assume the country is large with respect to all trading partners, so that the import supply curve originating from any trading partner is upward-sloping. To calculate the optimal exemption share θ_c for each country, we postulate the following:

(i) $\frac{dp^W}{d\theta_c} > 0$, a higher exemption share increases the world market price (this corresponds to assuming no Lerner paradox).

(ii) $\frac{dp}{d\theta_c} < 0$, a higher exemption share lowers the domestic price (this corresponds to the absence of the Metzler paradox).

(iii) $\frac{\partial M_c}{\partial \theta_c} > 0$, a higher exemption share for country c's product increases country c imports, holding the domestic price constant, due to substitution towards the tariff-exempt imports.

(iv) $\frac{\partial M_b}{\partial p} < 0 \ \forall b \neq c$, a higher domestic price reduces equilibrium imports from any given country $b \neq c$.

(v) $\frac{dr_c}{d\theta_c} > 0$, the rate of return for FDI in country c increases with the exemption share. Here, $\frac{dr_c}{d\theta_c}$ is a total derivative and includes the indirect effect via the world market price change.

(vi) $\frac{dr_b}{d\theta_c} < 0 \ \forall b \neq c$, the rate of return for FDI in countries other than c is negatively affected by an increase in the exemption share for c. This assumption implies that the positive effect of a higher world market price on rental returns in countries other than c is not sufficient to offset the negative effect of substituting away from imports originating from countries other than c.

We now calculate the first-order condition of maximizing (3.1) by choosing θ_c, assuming that the second-order condition of welfare maximization holds. We obtain:

$$\begin{aligned}\theta_c\tau[p^W(\frac{\partial M_c}{\partial p}\frac{dp}{d\theta_c} + \frac{\partial M_c}{\partial \theta_c}) + M_c\frac{dp^W}{d\theta_c}] \\ = [\tau\frac{dp^W}{d\theta_c} - \frac{dp}{d\theta_c}]M - \tau p^W M_c + (\alpha_d - 1)S\frac{dp}{d\theta_c} + \tau p^W\frac{dM}{dp}\frac{dp}{d\theta_c} \\ + \alpha_c\frac{dr_c^*}{d\theta_c}\hat{K}_c + \sum_{b\neq c}\alpha_b\frac{dr_b^*}{d\theta_c}\hat{K}_b. \quad (3.2)\end{aligned}$$

From the first-order condition, we can derive the comparative statics for our model. We conclude that the optimal exemption share θ_c is:

(i) increasing in total imports M,

(ii) decreasing in imports M_c from country c,

(iii) decreasing in domestic production S,

(iv) increasing in FDI $\hat{K}_c$ in country c,

(v) decreasing in FDI $\hat{K}_b$ in countries $b \neq c$.

The impact of an increase in the MFN tariff τ is in principle ambiguous. A higher (lower) MFN tariff will lead to a lower exemption share for country c's product if the substitution effect of a change in θ_c on imports is stronger (weaker) than the import-increasing effect of the lower domestic price.

These are the empirical predictions we take to the data.

4. Empirical Strategy

As in both Baier and Bergstrand (2004) and Magee (2003), we adopt a standard qualitative choice approach for the empirical estimation. The latent variable is interpreted here as representing the U.S. preference for offering preferential (zero-tariff) market access to a particular trading partner for a particular product in a particular year, and its associated observable choice variable measures whether or not preferential access is in fact offered. Let θ^*_{cjt} denote the latent value of offering preferential access for a given country c and product j at time t. Clearly, the latent variable need not be in the set $\{0,1\}$. First of all, considering the theory, it is quite reasonable to assume that θ^*_{cjt} lies inside the interval $(0,1)$. Moreover, it might even be negative (meaning that in principle, the U.S. would like to impose a tariff above MFN level) or above 1 (meaning that the U.S. would like to offer an import subsidy). Trade law requires, however, that preferences be binary; that is, the imports of product j from country c at time t must be either completely tariff-exempt or subject to the MFN rate.[7] Therefore, the observed tariff exemption θ_{cjt} is either 0 or 1, depending on whether welfare is higher when the MFN tariff is applied compared to when an exemption is granted or vice versa.

Although we cannot observe θ^*_{cjt} directly, we can observe many of its determinants. We define the econometric model:

$$\theta^*_{cjt} = \tilde{\alpha}_0 + \tilde{\alpha}_1 FDI_{cjt} + \tilde{\beta} \cdot X_{cjt} + \tilde{\gamma}_c + \tilde{\gamma}_j + \tilde{\gamma}_t + \tilde{\epsilon}_{cjt}, \tag{4.1}$$

where FDI_{cjt} is a measure of U.S. export-oriented investment in country c for producing product j at time t, X_{cjt} is a $k \times 1$ vector of other explanatory country-time, product-time,

[7] 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).

and country-product-time pair characteristics, $\tilde{\alpha}_0$ and $\tilde{\alpha}_1$ are scalar parameters, and $\tilde{\beta}$ is a $1 \times k$ vector of parameters. The parameters $\tilde{\gamma}_c$, $\tilde{\gamma}_j$, and $\tilde{\gamma}_t$ stand for country-, product-, and time-fixed effects, respectively. The remaining error term, $\tilde{\epsilon}_{cjt}$, represents unobserved heterogeneity in each country-product-time triple and is assumed to be independent of both X_{cjt} and FDI_{cjt}.[8] To the extent that the errors are correlated within countries or industries, country- and industry-fixed effects will correct for the Moulton problem. In the pooled sample we adopt for a robustness test, we cluster by both country and industry following the two-way clustering technique laid out in Cameron, Gelbach, and Miller (2006). (All of our specifications include year-fixed effects.) Clustering by both country and industry is substantially more conservative than in one dimension alone, and more so too than by country-industry pair (though we do cluster by country-industry pair as a robustness check in the panel specification).

Assuming that $\tilde{\epsilon}_{cjt}$ is normally distributed, the model in (4.1) should be estimated using Probit. The theory predicts that among otherwise identical country-product-time pairings, U.S. trading partner industries with more export-oriented FDI should have an increased likelihood of receiving preferential market access. Thus, the key theoretical prediction is that $\tilde{\alpha}_1 > 0$.

There are, however, several practical complications that call for modifications to our estimation strategy. First, several of our variables are not available at the product level j, but only at the industry level i, where each industry produces several products. In particular, while preferential market access is determined at the 8-digit HTS level, the investment data are available only at the more aggregated 4-digit NAICS level. For this

[8]One interpretation of the error term is that it captures unobserved noise in the legislative process by which preferential trade policies are determined. For instance, in the case of GSP preferences, the annual review by which policies are adjusted each year includes a series of petitions, public hearings, and a formal comment period, after which the executive branch (under the interagency Trade Practices Staff Committee) makes its decisions for the subsequent year. Other trade preference programs go through the U.S. House Ways and Means Committee or Congress as a whole, introducing further scope for unobserved political influence.

reason, the dependent latent variable is:

$$\theta^*_{cit} = \frac{\sum_{j\in i} M_{cjt}\theta^*_{cjt}}{\sum_{j\in i} M_{cjt}}, \tag{4.2}$$

where the product level trade weights $(\frac{M_{cjt}}{\sum_{j\in i} M_{cjt}})$ are from the year preceding the beginning of our sample period, and the observed industry level exemption share θ^*_{cit} then comes from the interval $[0, 1]$ rather than just taking on value 0 or 1.[9] Moreover, because t in our data is a year rather than just a point in time, even θ_{cjt} may take on intermediate values if the exemption status of a product switches within a year for any given country.

For both of these reasons – aggregation and part-year program eligibility – the dependent variable in our data must be treated as continuous. Our econometric model thus becomes:

$$\theta_{cit} = \alpha_0 + \alpha_1 FDI_{cit} + \beta \cdot X_{cit} + \gamma_c + \gamma_i + \gamma_t + \epsilon_{cit}, \tag{4.3}$$

where i stands for a 4-digit NAICS industry and t stands for years from 1997 to 2006.

The vector X_{cit} includes measures for U.S. domestic political pressure (U.S. domestic sales, payroll, total imports, import penetration, number of employees, and the log changes in U.S. employment and import penetration); the U.S. MFN ad-valorem tariff rate and an interaction term between the MFN tariff and an indicator for MFN tariffs above 1% for industry i and year t;[10] U.S. MNE sales from the rest of the world (other than country c) to the U.S. for industry i and year t; exports to the U.S. from country c in industry i and year t; and two gravity variables – GDP per capita and population for country c in year t. In the pooled sample, we also include country-specific, time-invariant characteristics for distance from the U.S. and indicator variables for whether

[9] In a robustness check, we consider the alternate non-trade weighted definition of $\theta^*_{cit} \equiv \frac{\sum_{j\in i} \theta^*_{cjt}}{\sum_{j\in i} 1}$ and reach qualitatively similar findings.

[10] Theoretically, we expect a negligible effect of very low MFN tariffs on the preferential tariff rate, since the foregone tariff revenue is very small. To account for this potential non-linearity, we include a dummy variable for MFN tariffs above 1% interacted with the MFN tariff. In the panel specification, this interaction with the dummy is necessary, as the latter is generally time invariant and would otherwise be absorbed in the industry-fixed effect.

the country was a communist or terrorist state during our ten-year sample period, as well as time invariant industry-level dummies for agriculture and textile sectors.

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 (4.3) 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- and country-fixed effects γ_i and γ_c by demeaning the data – but of course we must then ignore empirically the censoring process that generates the mass points for θ_{cit} at 0 and 1.[11] In a second version of the empirical strategy, we adopt the non-linear double censored Tobit model with country-, industry- and year-fixed effects included as dummy variables, but the need to achieve convergence limits both our choice of estimator and the set of control variables we can include.[12]

Finally, we address the potential simultaneity between export-oriented investment and preferential access by instrumenting for our FDI measure. Finding suitable excluded instruments that predict export-oriented FDI, but are at the same time uncorrelated with the error term, is in general quite challenging. Fortunately, the BEA-DIA data separates MNE sales data by destination. According to the theory, import oriented (horizontal) FDI should be independent of U.S. trade policy, but at the same time, it seems likely (and we confirm in the data) that import oriented and export-oriented FDI are positively correlated, presumably because both capture the attractiveness of the local market for foreign investors. Thus, we instrument for export-oriented sales to the U.S. with MNE

[11]Year-fixed effects are still removed by dummy variables. Because our data set is not balanced, we take care to demean the year dummies. We also correct the standard errors to account for the demeaning (to account for the difference in degrees of freedom); given the large number of observations, the standard error correction is negligible, however.

[12]To achieve convergence in the instrumented Tobit specification, we adopt the Newey (1987) efficient two step estimation and reduce the set of controls to the bare bones minimum called for in the model. (In the other baseline runs, we find a small set of gravity and political economy variables to be sensible additions as they prove statistically significant.)

sales to the *local market*.[13] In a just-identified specification of the model, local MNE sales is the only excluded instrument; we also include the square of local MNE sales as additional instrument in an over-identified version of the model to test for instrument validity.

5. Data

Preferential Market Access. The dependent variable of interest is preferential market access by industry, country, and year to the U.S. market. There are two ways to construct the preferential access variable using slightly different data sources; we consider both definitions to evaluate the robustness of our findings, and thus to ensure that our results rest on meaningful economic factors rather than specific variable definitions.

Our first data source for preferential market access comes from the U.S. Trade Representative harmonized tariff schedules (HTS-US). Imposed tariff rates and the relevant indicators for preferential program eligibility are reported at the 8-digit HTS level by country, industry, and year.[14] Thus, one way to define the preferential treatment variable is so that $\theta_{cjt} = 1$ if the country-product pair is eligible for a special rate code in year t, and 0 otherwise.[15] When we aggregate to the 4-digit NAICS level (necessary to concord the preferential tariff data to the investment data), we construct both historic trade

[13] One might reasonably raise the concern that FDI 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, however, which argues against the substitutes story. Moreover, the broader 4-digit NAICS industry categories allow a host of firms selling a range of products within each industry. In general, the firms and products selling to the local market are distinct from those selling to the U.S., so there is good reason to believe that local sales will be independent of U.S. trade preferences. (On average, less than 25% of firms in our sample sell both locally and to the U.S.) Indeed, in a simple pooled OLS version of the model, we found that export-oriented MNE sales were positively and significantly correlated with U.S. trade preferences at the 1% level, while MNE sales to the local market were uncorrelated.

[14] As of January 1, 2006, the HTS-US schedule included 19 preferential treatment codes for GSP (and its subcategories), regional agreements, etc.

[15] We code a country-product-year observation as preference eligible if it is eligible for more than one quarter of the given calendar year.

weighted[16] and straight (unweighted) averages across the relevant subcategories. We view the former as the more appropriate measure, as it captures the empty promise of preferences for goods that are not produced by a beneficiary country (mangoes from Iceland, semiconductors from Afghanistan, etc.). In the tables, we label the trade weighted eligibility measure of duty free access under any preference program *El Any*. We report robustness results for the unweighted eligibility measures in the footnotes of the Results section; the results are qualitatively unchanged by removing the trade weights.

While this simple definition of the preferential treatment variable is appealing in its parsimony, one can easily challenge the definition on the grounds that even when preferential eligibility is indicated by the HTS-US, preferential treatment may be afforded to only a subset of the imports in question. Partial-year program eligibility is a key concern, as many program changes (including virtually all changes under GSP) are effective July 1st (or idiosyncratically), rather than January 1st of a calendar year. Moreover, GSP preferences can be (and often are) limited by additional "competitive need limitations" (CNLs), which offer duty-free treatment only until a certain level of exports is reached. Finally, highly restrictive rules of origin restrictions or other bureaucratic costs under some programs may make *de jure* preference eligibility useless in practice. A preference measure based on countries' actual usage of the programs can capture such otherwise unobserved limitations to program use in practice.

With these caveats in mind, we define a second form of the dependent variable 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.[17] We use this information

[16]Time invariant trade weights are constructed using 1997 trade flows, the year immediately preceding the first year in our sample.

[17]For instance, when a product enters the U.S. under a GSP eligibility clause, it receives special tariff code A (or A*, A+ depending on the particular sub-classification of GSP eligibility). When a product enters the U.S. duty-free under a free trade agreement, then an agreement-specific code is entered (for instance "MX" for Mexican products entering under the NAFTA or "R" for products entering under the Caribbean Basin Trade Partnership). These codes match those used by the USTR.

to construct our baseline measure of the dependent variable, *Any Pref Share*, so that θ_{cjt} is the (exact) share of country c exports of product j in year t that enter U.S. Customs claiming duty free access under any preferential program code. Note that this version of the dependent variable based on U.S. Customs data offers the additional feature that it does not require an ad-hoc weighting scheme to aggregate to 4-digit NAICS.[18]

In addition to these two measures of U.S. preferential market access under any and all preference programs, we create three alternative dependent variables: two measures of duty free market access under GSP only – one eligibility based, *El GSP*, another based on Customs data and actual GSP use, *GSP Share*, and a third alternative for a robustness check, *Non-GSP Share*, the rate at which a country's exports claim any preferential market access code other than GSP. There are several reasons that we find GSP preferences of particular interest. The first is simply that the GSP program comprises roughly two-thirds of country-industry-year observations with preferential market access in the U.S.[19] At the same time, the mechanics of the GSP program have important institutional differences from regional or other preferential programs. There is a formal process of annual reviews of GSP eligibility in which domestic and foreign firms, foreign

[18]It should be noted that some of the deviation between the Customs-use and eligibility-based measures is caused by foreign exporters failing to (or choosing not to) claim preferential access when eligible. (See recent work by Hakobyan (2010) on underutilization of GSP preferences.) The relevant concern for our study would be if MNE foreign affiliates are more able to *use* preferential trade programs than are commensurate locally owned firms (which, to be fair, is not obvious – while U.S. MNE affiliates may find it easier to file the paperwork necessary to get preferential market access, their disproportionate use of sourced inputs may make rules of origin harder to satisfy). The robustness of our results to either preference variable definition – program use, which could potentially be contaminated by differential uptake of preference program usage by MNE affiliates (but is otherwise a more precise measure), and eligibility, which is clearly free of any such concern (but is also a less perfect measure given aggregation and related issues outlined earlier) – suggests that this discrepancy is of minimal consequence in the context of our study.

[19]Weighting by trade volume, however, GSP is predictably a smaller share; total U.S. imports under GSP comprise 7% all preferential imports.

governments, labor unions, and other interest groups can, and regularly do, take part.[20] Year to year, GSP eligibility can be revoked on a discretionary product level basis due to human rights, labor, or intellectual property violations, and is regularly limited by binding competitive need limitations that cap imports from the most productive developing countries. Finally, from an econometric perspective, the GSP-based preference measure is largely immune from concerns over reciprocal trade policy implications or bilateral investment protections that may arise under Article XXIV style free trade deals.

It is worth spending a moment to examine the extent of variation in our dependent variable, however it is defined. From a literal reading of GATT rules, one would expect limited discretion in the extent to which preferential market access can vary across countries, industries, and particularly country-industry specific pairs. Article XXIV, which governs preferential free trade or regional trade agreements, specifies that among signatory countries duty free access should apply to "virtually all" products. In principle, then, we would expect that country-fixed effects (or even more conservatively, country-year fixed effects) would explain virtually all of the observed variation in trade preferences under regional or bilateral free trade deals. The Enabling Clause, which authorizes GSP programs, allows more discretion at the industry level – certain industries may be excluded from GSP eligibility entirely – but these industry exclusions must be uniform across GSP beneficiary countries, effectively putting *generalized* in the Generalized System of Preferences.[21] In principle, we then would expect that together, country- and industry-fixed effects would account for virtually all of the observed variation in our dependent variable.

Of course, it is well understood that there is some leeway in incorporating exemptions and exclusions both during the initial negotiation of preferential agreements and through

[20]The GSP annual review process as well as CNL implementation, waivers, and waiver exemptions, are determined by the U.S. Trade Representative in a cabinet subcommittee through a process of public hearings and petitions, with considerable discretion ultimately left to the executive branch.

[21]The U.S. GSP includes a two-tiered system to allow enhanced market access for "least developed beneficiary countries," but as will be clear from both the following figures and our empirical results, the two-branch system is not responsible for our results.

subsequent formal and informal review processes thereafter. To convince the reader that there is, in fact, sufficient variation left to explain after including country-, industry-, and year-fixed effects, we offer the following simple plots. On the vertical axis, we plot the residual of our dependent variable, preferential market access after controlling for country-, industry-, and year-fixed effects. On the horizontal axis, we plot country per capita GDP. Each plot is for a different definition of the dependent variable: *Any Pref Share* and *GSP Share* in the top row, and *El Any* and *El GSP* in the bottom row. For the two GSP-specific measures, we include data only for the set of countries that is potentially GSP eligible. In all four plots, we see quite clearly that there is considerable variation in every measure of preferential market access and that the degree of variation is higher among the U.S.' less developed (low GDP per capita) trading partners.

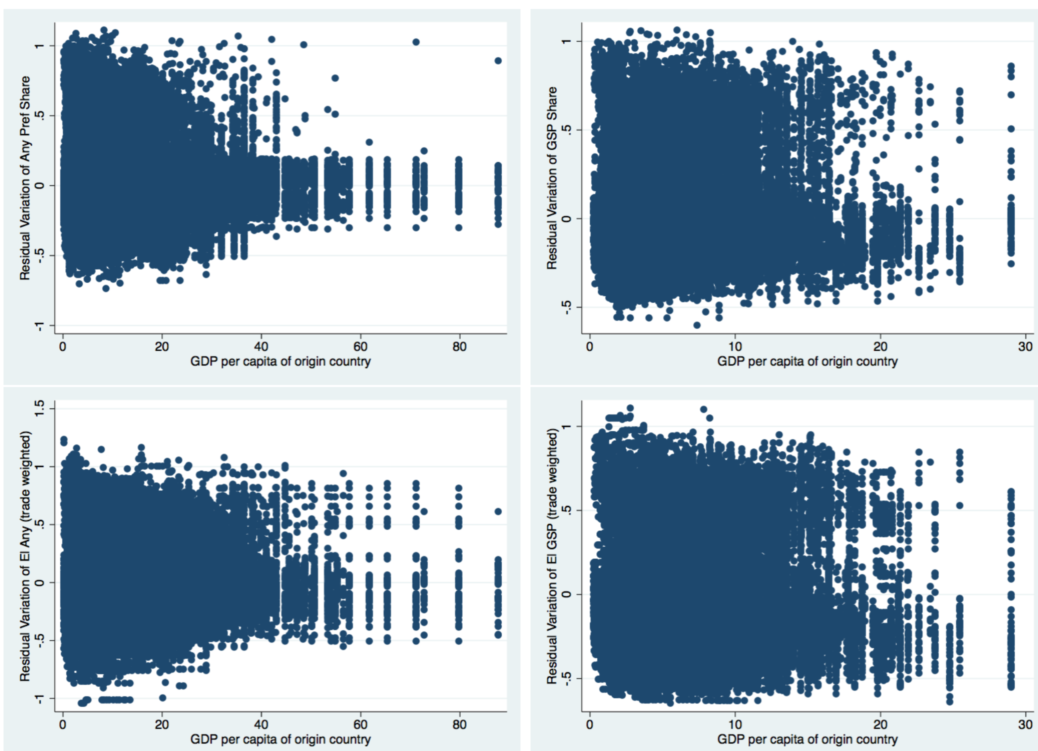

FIGURE 1. Dependent Variable Residual Variation

Multinational Affiliate Activity. Our foreign investment data are from the U.S. Bureau of Economic Analysis data on U.S. direct investment abroad. The BEA data consist of 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.[22] The 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 (MNC) operations.) Industries are coded at the 4-digit NAICS level, so that each 4-digit NAICS industry includes multiple 8-digit HTS product lines.[23, 24] For the purpose of testing equation (4.3), we define FDI_{cit} as multinational affiliate sales to the U.S., prorated by the percentage of U.S. ownership.[25] Recalling the theory, the influence of export-oriented investment on U.S. tariff preferences depends on the sensitivity of FDI returns to θ; i.e. $\frac{dr^*_{cjt}}{d\theta_{cjt}}\hat{K}_{cjt}$. We argue that sales are the most accurate measure of the sensitivity of profits to changes in preferences (via the selling price); by Hotelling's lemma, recall that the derivative of the profit function with respect to output price p is supply ($\frac{\partial \pi(p,\vec{w})}{\partial p} = y(p,\vec{w})$) if factor prices $\vec{w}$ are fixed. Indeed, if foreign investors are the residual claimants of affiliate profits and view local wages as given, then our sales measure is *exactly* the derivative theory suggests. For our instruments, we use MNE sales to the local market (and its square), also from the BEA data.

[22]We necessarily restrict our sample to majority owned foreign affiliates (MOFAs), however, as only MOFAs report sales disaggregated by destination. 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.

[23]In a few instances, a given 8-digit HTS code concords to multiple 4-digit NAICS, in which case we divide the HTS8 import data evenly among the relevant NAICS codes.

[24]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 the standard 6-digit NAICS.

[25]Most foreign affiliates in our sample are wholly U.S. owned, and the prorating does not influence our results.

Control Variables. Finally, we include a number of control variables at the country, industry, country-year, industry-year, and country-industry-year levels. When we include country- and industry-fixed effects, of course, the time-invariant country and industry level controls (such as distance to the U.S. and indicator variables for textile or agricultural industries) are dropped. Table 1 summarizes the variables that we include in our data set and their sources.

In principle, our data set would have 204,160 observations: 232 countries by 88 industries for 10 years (1997-2006). In practice, however, our sample is smaller. We have data for a subset of 184 countries and 80 industries over the ten-year sample. When we lag the independent variables for one year (as seems most appropriate given the time needed for policy to change) and include variables for the log change in U.S. employment or import penetration, our data are further reduced to a 9-year panel beginning in 1998.[26] Of these remaining data, another limitation arises. When we report the preference variable as the share of imports entering the U.S. with preferential treatment, we lose all observations for which U.S. imports (the denominator of the ratio) are zero, leaving 68,130 observations. To make our baseline results consistent across specifications, we use this smallest data set for all specifications of the model, even in cases where we have access to more observations (i.e. for the eligibility-based definitions of the preference variables, which need not preclude observations for which the U.S. import volume was zero). We conducted robustness checks to compare the largest possible sample size with our estimates from the smallest commensurate sample size, and the findings are consistent throughout, though of course the point estimates vary somewhat (non-systematically) and the standard errors are generally slightly larger with the smaller sample. Table 2 reports descriptive statistics for the baseline data set with 68130 observations. Table 3 provides the same statistics for the data set with only de jure GSP eligible countries

[26]While some preference programs are reviewed only on an ad hoc basis, others, like the GSP, have a formal annual review process for petition, study (by the USITC), and ultimate implementation by USTR on a regular yearly schedule.

included, which contains 42849 observations. An overview of the NAICS4 industries and the countries included in our sample is given in Table 4.

6. Results

Table 5 reports the estimation results when the dependent variable is 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. Country- and industry-specific fixed effects are removed by demeaning, and time-fixed effects are incorporated as year dummies.[27] The table reports coefficient estimates for only the variables of major interest. Explanatory variables included but not reported are U.S. sales squared, U.S. industry payroll and number of establishments, U.S. industry employment, the log change in U.S. import penetration from the previous year, and the unconditional U.S. MFN ad-valorem equivalent.[28] The first column reports the results of the just-identified IV model. With a few exceptions, the variable coefficients have the expected sign and are statistically significant at the 10% level or better.[29] In particular, higher overall imports lead to an increase in preferential treatment, as predicted by theory. The estimate for U.S. domestic sales also takes the predicted negative sign, and is significant at the 1% level. The coefficient for the MFN average tariff rate is positive and also significant at the 1% level; recall that the theoretical prediction is ambiguous. U.S. MNE sales to the rest of the world and country-industry-year exports to the U.S. appear to have little effect on the rate of preferential market access, with a p-value of .40 and .29, respectively. We also find that countries with lower GDP per capita receive ceteris paribus a greater rate of preferential market access, likely under the auspices of the GSP, although this effect is

[27]Standard errors are corrected accordingly.

[28]The table reports the estimate for the MFN average interacted with a dummy variable for MFN rates above 1%; we include both MFN terms to treat separately those industries with negligible MFN tariffs (less than 1%).

[29]Of the other included variables, the following are statistically significant at the 10% level: U.S. establishments (thousands) [.0002 ($p = .02$)], U.S. employees (millions) [$-.08$ ($p = .02$)], the MFN tariff [-3.05 ($p < .001$)] and U.S. sales squared (bUSD) [2.3 ($p < .001$)].

not statistically significant. Interestingly, there is a much stronger relationship between an exporting country's population and the granted preferences: quite apparently, less populous countries receive ceteris paribus higher trade preferences. Last, as one might expect, healthier U.S. industries with growing employment and lower import penetration ratios tend to offer more generous preferential access; struggling U.S. industries garner more protection.

Most importantly, an increase in multinational sales to the U.S. of $1 billion implies a statistically significant (at the 1% level) increase in the exemption share by 3.5 percentage points. Conditional on having any FDI initially,[30] a one standard deviation increase in the mean level of export-oriented MNE sales for a given country and industry then implies a more than 3.25 percentage point jump in the rate of preferential imports; this in turn translates to a 16.3 percent increase in the rate of preferential access relative to the sample average of .20. Finally, note that the excluded instrument performs well in terms of explaining the possibly endogenous variable "MNE sales to U.S.", as is apparent from the first-stage F-statistic of 27.45.[31] Because our model is just identified, however, we cannot yet test for instrument validity.

In the next three columns, we reestimate the model to test for instrument validity and specification robustness. Each variation includes a second excluded instrument, squared local MNE sales, and uses two-stage least squares, two-step GMM, and GMM with clustering at the country-industry level, respectively. The J-statistic p-values consistently exceed 0.1 by a substantial margin, passing at least this simple test for instrument validity. The coefficient estimates for MNE sales to the U.S. remain quantitatively consistent across the various specifications, increasing slightly compared to the just identified model in the two GMM specifications and hovering around a 3.4 to 3.9 percentage point increase in the rate of preferential market access for each $1 billion in MNE sales to the U.S.

[30]Among countries with any U.S. FDI, the mean FDI sales to the U.S. (prorated by U.S. ownership) is .087 billion USD with a standard deviation of .929 billion USD.

[31]The first stage coefficient for the instrument (MNE local sales) is .21 ($p < .001$).

In Table 6, we report the results for a similar set of estimations when the dependent variable is the percentage of imports for which a GSP exemption was claimed at customs. Again, the coefficient estimates are generally consistent across the various specifications of the model. One change is that U.S. total imports, which were positively correlated with the all preferences share, are negatively correlated with GSP preferences. Also, the effect of country-industry-year exports of a country on GSP preferences is of similar magnitude as before, but is now statistically significant in most cases. Both findings are broadly consistent with binding competitive need limitations, which limit the extent to which successful exporters can claim duty free access under GSP. In contrast to the previous dependent variable, *Any Pref Share*, the GSP-specific measure of trade preferences *GSP Share* is responsive to per capita country GDP, with poorer countries receiving higher GSP preferences as should naturally be the case given that rich countries are not GSP eligible.[32] At the same time, the pure size effect, i.e., smaller countries garner higher preferences, is still present.

Turning to the key estimate of interest, we again find that the effect of U.S. multinational goods sales to the U.S. continues to be statistically significant (except for the clustered specification), but the estimate is somewhat smaller: an increase in multinational sales to the U.S. of $1 billion is associated with a statistically significant (at the 1% level) increase in the GSP program exemption share of slightly more than two percentage points. The implication seems to be that while GSP preferences appear to be influenced by offshoring MNE activity, the per-dollar effect of FDI on GSP preferences is roughly two thirds that for all preference programs in general. This conclusion may be hasty, however, once we recognize that a substantial number of countries included in this sample are automatically excluded from the GSP program according to U.S. law, and the relevant coefficient estimates thus may be biased downward.[33] Also of concern

[32]Moreover, the poorest countries are eligible for enhanced GSP market access when accorded LDBC (Least Developed Beneficiary Country) status.

[33]An additional complication to the GSP results relative to the findings for all preference programs together is that many GSP eligible countries (and certainly those most likely to export to the U.S.) are also beneficiaries of other more generous unilateral agreements: CBI, AGOA, and the Andean

is the very small p-value for the Hansen's J test of instrument validity – it seems that for the GSP variable in the full sample, the instrument is likely correlated with the second stage error term. We speculate that MNE activity of any sort (including our instrument – MNE sales to the local economy) is positively correlated with a country's level of development, and thus (inversely) to a country's potential eligibility for the GSP program. (Most horizontal FDI goes to wealthy countries that by law are ineligible for GSP preferences.)

In Table 7 therefore, we further restrict the sample to exclude countries that are *de jure* GSP ineligible under the 1974 Trade Act.[34] In the reduced sample, 91% of observations are for currently GSP eligible countries, opposed to 57% in the full sample. We report results for both definitions of the dependent variable: the share of U.S. imports that receive preferential access under the GSP program only (*GSP Share* in columns (1) and (2)) and, as a robustness check, the *Any Pref Share* version of the dependent variable in columns (3) and (4).

The effect of removing the *de jure* GSP ineligible countries is striking: the coefficient estimates for the impact of MNE sales on U.S. trade preferences increase by more than an order of magnitude across the board. In interpreting the coefficient estimates, it is important to keep in mind that the average level of MNE sales to the U.S. is much lower

trade pact. GSP usage may be lower for these countries than their otherwise identical counterparts simply because they have more generous market access (primarily through less restrictive rules of origin) through other programs. To the extent that FDI also induces more generous non-GSP treatment, our GSP results will be biased downward.

[34]Countries are immediately 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, as some otherwise *de jure* GSP eligible OPEC members are consistently excluded (such as Iran and Libya), while others are not (such as Algeria, Angola, Ecuador, Iraq, Nigeria, and Venezuela for 2009). The 'terrorism' criterion is even less transparent. We therefore restrict the sample to exclude countries under only the first three criteria.

in the reduced sample. Thus, conditional on having any FDI,[35] a one standard deviation increase in MNE sales to the U.S. is associated with roughly a 14.8 percentage point expansion of the rate of GSP access – a 78 percent increase relative to the (reduced) sample average rate of GSP preferential access of 19%. Perhaps surprisingly, we now see that preferences writ large – measured by the *Any Pref Share* variable – are somewhat *less* responsive to MNE sales than are GSP preferences specifically. Comparing columns (1) versus (3) or (2) versus (4), the dollar for dollar impact of MNE sales to the U.S. is larger for GSP preferences than for trade preferences overall.

The instruments are somewhat weaker with this restricted sample, presumably because there are fewer industries among the developing countries with MNE operations in both the export-oriented and import-competing (local sales) sectors. At the same time, instrument validity appears to be less of a concern given the J-statistic p-values of more than 90%. We are again comforted to find that the estimation technique (GMM or two stage least squares) has little influence on the result.

Scanning the remaining coefficient estimates in Table 7, we find them broadly consistent with the earlier estimates from the full data set in Tables 5 and 6. One noteworthy exception is the effect of country population on trade preferences. Whereas in the full sample, smaller countries reaped higher trade preferences, the effect is opposite once we only consider de jure GSP-eligible countries: *within the set of GSP eligible countries*, more populous countries receive higher trade preferences ceteris paribus. All other coefficient estimates take the same sign and ball-park statistical significance as in the commensurate runs for the full data set. That said, we again find the coefficient estimates for this reduced data set are typically larger (in absolute value) by as much as an order of magnitude than the corresponding estimates from the full data set, suggesting that U.S. trade preferences for developing countries may be more sensitive to the underlying economic and political environment than are U.S. trade preferences to the world at large. A priori, one might have anticipated that trade preferences for poorer

[35]Among *de jure* GSP eligible countries with any U.S. FDI, the mean FDI sales to the U.S. (prorated by U.S. ownership) is .032 billion USD with a standard deviation of .289 billion USD.

countries would be less subject to domestic U.S. influences such as import penetration, change in the number of employees, and U.S. domestic sales, but we find no support for such a view.

We now turn in Table 8 to an IV Tobit specification of the model 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 now include country-, industry-, and year-fixed effects through dummy variables. Due to computational limitations in the estimation of the IV Tobit model with three dimensions of fixed effects, we both reduce the set of controls to include only those explicitly specified by the model and resort to the Newey (1987) two step estimation to achieve convergence.

The first two columns of Table 8 report results for the full sample for each variation of the dependent variable (share of imports entering the U.S. duty free under (*i*) any preference program or (*ii*) GSP specifically) while the second two columns repeat the same procedure for the reduced data set, eliminating *de jure* GSP ineligible countries. The results are qualitatively similar to the results of the IV panel estimation, in that we find a strong positive effect of MNE sales to the U.S. on preferences. For the *Any Pref Share* variable in column (1), the effect of MNE sales to the U.S. is roughly double in size compared to the IV panel specification.[36] Unsurprisingly, given the implicitly censored nature of the dependent variable *GSP Share* in the full sample, we find that the IV Tobit specification in column (2) yields a much larger coefficient estimate for MNE sales to the U.S. relative to the previous linear specification in Table 6. Reducing the sample to include only the de jure GSP-eligible countries in columns (3) and (4) leads to an increased effect of FDI on preferences, consistent with our linear model findings in Table 7. In particular, the estimated effect on GSP preferences is almost the same in the IV panel and the IV Tobit specification, whereas the effect on the all preference share is smaller and loses significance in the IV Tobit compared to the IV panel specification. Finally, note that most of the censoring is at zero (55% of the observations for the

[36]The coefficient on *MNE Sales* is .066*** in Table 8 col. (1) versus .035*** in Table 5 col. (1).

all preference share variable) 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 the bias relative to the linear IV results.

Summarizing our results thus far, we draw three broad conclusions. First, the empirical results are qualitatively consistent with the model, and our instrument FDI sales to the local market performs strongly in the full sample and only somewhat less so in the reduced sample. The key finding, of course, is that U.S. multinationals' U.S.-destined sales seem to increase the rate at which exports from the MNE host country and industry will be afforded preferential market access. Our empirical specifications control for the likely endogeneity of export-oriented sales and incorporate country-, industry-, and year-fixed effects. Under the linear panel IV specification, a one standard deviation (roughly $1 billion) increase in MNE sales to the U.S. is associated with roughly a three and a half percentage point increase in preferential tariff exemption for the sample of all countries.

Second, while in the full sample the GSP program appears to be less responsive, dollar-for-dollar, than preference programs in general, when we restrict the sample to exclude *de jure* GSP ineligible countries, we find the estimates for both types of preferences to rise by more than an order of magnitude 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. Recognizing the lower average level of MNE sales in developing countries, the impact of a one standard deviation (roughly $290 million) increase in MNE sales to the U.S. is a 14.8 percentage point increase in the rate of GSP access to the U.S. – nearly a 78 percent increase in the percentage of exports afforded GSP access relative to the sample mean among GSP eligible countries.

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. The effect of a $1 billion USD increase in MNE sales back to the U.S. entails a remarkable

6.6 percentage point increase of the rate of any preferences in the full sample, and a 58.8 percentage point increase in the rate of GSP preferences under the reduced sample.

7. Robustness Checks

We conduct a battery of robustness checks. First, to double-check the instrumentation strategy, we run a simultaneous equation GMM specification of the baseline model, delivering estimates for *both* directions of causality: the influence of MNE sales on trade preferences (which we find qualitatively and quantitatively consistent with our existing results) and also the effect of U.S. trade preferences on MNE activity (also positive and statistically significant, as expected). In a second step, we examine the robustness of our model, reporting results for several alternative model specifications and alternative definitions of the dependent variable based on program eligibility. A third set of tests looks even further into within-group variation, running the baseline (panel IV) specification once each with three different sets of pairwise fixed effects (country-industry, industry-year, and country-year). Finally, we work through a short list of potential data concerns, confirming that our basic findings are robust to alternative aggregation weights, exclusion of non-BEA-benchmark years, and inclusion of interaction terms for time-invariant country and industry control variables.

Simultaneous Equations. We first consider a system of simultaneous equations where we model both the (already much examined) effect of export-oriented FDI on trade preferences and also the reverse causal effect of trade preferences on FDI. We use three-stage least squares to estimate and test this system of simultaneous equations. The preference equation contains the same variables as in the panel IV specification. We choose to report a slightly smaller subset of variable coefficients in the interest of space, however; the two additional unreported variables country GDP per capita and U.S. import penetration are generally no longer statistically significant. In the FDI equation, we use our earlier instruments, local MNE sales and its square, total imports from country c in industry i at time t, the average MFN tariff and the MFN tariff interacted with a dummy for values greater than 1%, country GDP per capita, population, year dummies, and trade preferences as explanatory variables. Both equations are overidentified in this specification.

The top half of Table 9 reports estimates for the FDI equation. Despite our admittedly ad-hoc specification of the FDI equation, the parameter estimates show a clear statistical relationship between the chosen explanatory variables and our FDI measure. The first row of results from the simultaneous equation specification provides strong evidence that causality runs both ways. In the full sample, a one percentage point increase in the rate of preferential market access (captured in the first column, *Any Pref Share*) increases MNE sales to the U.S. by .59 billion dollars; a one percentage point increase in the rate of GSP access (*GSP Share*) share variable of 1 percentage point increases MNE sales to the U.S. by nearly 1.8 billion dollars. As in the first stage of our earlier IV specifications, MNE sales to the U.S. are closely correlated with our instrument, MNE sales in the (foreign) local market, though the relationship appears to be weakly convex in the full sample and concave in the reduced sample.[37] Scanning further down the list of explanatory variables in the FDI equation, we find that total country-industry exports to the U.S. go hand in hand with more MNE sales to the U.S. Finally, we notice in passing that U.S. firms seem ceteris paribus to invest more in less populous countries, both in the overall sample and (perhaps more surprising) among *de jure* GSP eligible countries in the reduced sample.

Our main interest in the paper lies of course in the preference equation, the results of which can be found in the bottom half of Table 9. There, we see that the 3SLS estimates for the influence of MNE sales to the U.S. on preferential market access are quite similar to the IV panel results. For the full sample, the coefficient on MNE sales is basically identical with values of .036 for all preferences and .022 for GSP preferences. For the *de jure* GSP eligible country sample, just as for the panel IV specification, the influence of FDI on trade preferences increases. The coefficient estimates equal .343 for

[37]In both samples, the correlation between MNE sales to the U.S. and local MNE sales is positive and significant at the one percent level; in the full sample, the square of local sales is also positively correlated with FDI (though quite small), whereas the correlation for the quadratic local sales measure is negative in the reduced sample.

all preferences and .426 for GSP preferences and are statistically significant at the 1% level.

Interestingly, when we compare the results for the full and reduced samples in Table 9, we see that the estimated reverse causal effect of trade preferences on FDI becomes considerably smaller for the reduced sample (an estimated .22 billion dollars MNE sales increase for a one percentage point increase in the GSP preference variable; the effect of a change in all preferences statistically disappears) just as the reverse effect of U.S. MNE sales on trade preferences increases in magnitude. In other words, while causality clearly runs both ways in the full sample, among the set of *de jure* GSP-eligible countries, the causality running from FDI to preferences becomes stronger whereas the reverse causality nexus appears to lose some of its bite.

Alternative Specifications and Dependent Variables. Although we find the most sensible modeling specification to be the instrumented panel versions presented in the previous section, we now include for comparison an IV version of the model on cross-sectionally pooled data (without country- and industry-fixed effects) and two reduced form (no IV) linear versions of the model – one panel and the other pooled. Both pooled specifications include year-fixed effects and are corrected for two-way clustering at the country and industry levels.[38]

As a robustness check on our baseline definition of the dependent variable, we consider an alternative definition of trade preferences based on official program eligibility (rather than actual trade flows by program claimed, which is our baseline preference '-*Share*' variable definition), which we aggregate to 4-digit NAICS using historical (1997) trade volumes.[39] For completeness, we also include a set of regressions for the influence of MNE activity on non-GSP preferences (the share of U.S. imports that claim duty free access under any program *other* than GSP).

[38]We report significance for both unclustered heteroscedasticity-robust and two-way clustered standard errors, simply to note the influence of clustering for the interested reader; accounting for multi-dimensional clustering is clearly important.

[39]We also test unweighted aggregations of the preference eligibility variables and find the estimates little changed.

We report the results for each of these alternatives in Table 10 for the full sample and in Table 11 for the reduced sample. For ease of reference, we also include the already reported baseline results from the linear panel IV, IV Tobit, and 3SLS specifications. For brevity, the tables list only the coefficient estimate of interest – MNE sales to the U.S. – for each specification. Each row represents a given model specification, while each column features a different definition of the dependent variable. Thus, 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. Comparing results across the two tables, we again suggest focusing on the full sample results for the estimates of all preference programs in general or non-GSP preferences, and on the reduced sample results for GSP programs specifically.

Reviewing the pair of 'Alternative Specifications' Tables 10 and 11, 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, focusing on the baseline panel IV specification in the first row. Juxtaposing columns 1 and 3 for all preference programs and columns 2 and 4 for GSP, we see that the estimates are qualitatively similar across the two dependent variable definitions – actual preference program usage from U.S. Customs data (the *-Share* variables in columns 1 and 2) versus program eligibility from the USTR (the historical trade-weighted eligibility *El-* measures in columns 3 and 4).[40] The point estimates

[40]As an additional robustness check, we confirm that the trade weights used to aggregate program eligibility to 4-digit NAICS are not responsible for our findings in columns (3) and (4). In the baseline panel IV specification (top row), we find that the key coefficient estimates for *unweighted* versions of the dependent variables *El Any* and *El GSP* are slightly smaller than for the trade-weighted variables, but neither positivity nor significance are affected by dropping the trade weights: for *El Any*, the key coefficient estimate for the trade-weighted version is .017**, while that for the unweighted version is .013**; for *El GSP*, the key estimate for the trade weighted version of the dependent variable is .004, compared to .002 for the unweighted version. Removing the trade weights has a similar effect in the reduced sample: for the panel IV specification in Table 11, the key coefficient estimates in columns (3) and (4) decline from .46*** to .39*** and from .51*** to .43***, respectively.

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 much more closely in the reduced sample. Finally, note that the *GSP Share* and *Non-GSP Share* estimates sum to the *Any Pref Share* estimate values (as one would expect) and that the *GSP Share* point estimate is almost always higher than the *Non-GSP Share* equivalent, suggesting that GSP preferences are generally more responsive to vertical MNE activity than are non-GSP preferences.[41]

Comparing instead across the different model specifications in each row, we note the instrumental variables results to be quantitatively very similar across specifications in the full sample (Table 10), whereas in the reduced sample (Table 11) the Pooled IV model (which removes country- and industry-fixed effects) produces considerably higher estimates than the Panel IV model. The OLS specifications yield smaller estimates, whether or not country- and industry-fixed effects are included (recall that year dummies are universally applied), but they are typically also positive and significant.

To summarize, while the coefficient estimates vary, they are consistently positive and significant for all but a handful of the numerous variants of the model reported in the upper section of Tables 10 and 11.

Pairwise Fixed Effects. Pushing the data in a different direction, we explore within-group variation more deeply by including pairwise fixed effects in the bottom part of Tables 10 and 11. (Of course, doing so reduces within-group variation considerably, so estimates are predictably less precise.) We reran the baseline panel IV model three more times: first with country-industry pair fixed effects, time dummies, and the MNE U.S. sales variable; again with industry-year pair fixed effects, country dummies, and the MNE U.S. sales variable; and finally with country-year pair fixed effects, industry dummies, and the MNE U.S. sales variable. In each instance, we still find that the

[41]In another robustness check, we created a second sub-sample of *de jure* GSP *ineligible* (primarily wealthy) countries, and examined the influence of U.S.-bound MNE sales on Non-GSP preferences. The baseline coefficient estimate was $.013^{***}(p = .01)$ in the linear panel IV model; for the alternative model specifications, the key coefficient estimate ranged from $.01^{***}(p < .001)$ in the Panel OLS version of the model to $.04^{***}(p < .001)$ in the Pooled IV specification.

coefficient of our key explanatory variable of interest, MNE sales to the U.S., is typically positive and significant. When removing joint country-industry fixed effects (so that identification comes only from intertemporal variation over nine years), the estimates usually drop by a factor 10 in the full sample compared to the panel IV baseline model, whereas the other joint fixed effect specifications generally yield coefficient estimates that are quantitatively very similar to the baseline panel IV specification in the first row. For the reduced sample, the estimates are only slightly smaller than in the panel IV baseline case and more homogeneous across the different pairwise fixed effect specifications, even in the case of country-industry pairwise fixed effects.

Additional Robustness Checks. For brevity, we exclude here the final set of robustness checks, but all are readily available upon request. In these, we restrict the sample to only benchmark years for the BEA data (1999, 2004) to address potential concerns over the robustness of non-benchmark year BEA data; we introduce interaction terms for country- and industry-level time invariant control variables in the panel IV specification; we define the 4-digit NAICS aggregate MFN variable as an *unweighted* average of its sub-aggregates (in contrast to the trade weighted version used in the benchmark model), and we measure program eligibility using *unweighted* aggregation to 4-digit NAICS. For each of these variants, our baseline results remain qualitatively unchanged, though of course the point estimates vary somewhat and standard errors increase substantially for some of the smaller sample sizes.

8. 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 uniquely 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 find that within a given (4-digit NAICS) industry in a given country and year, each $1 billion in U.S. foreign affiliate exports to the U.S. is associated with roughly a 3.5 percentage point increase in the rate of preferential duty free access from all preferential programs combined. Thus, a one standard deviation ($970 million) increase in MNE export-oriented sales to the U.S. would induce roughly a seventeen percent (3.4 percentage point) increase in market access relative to the mean. Restricting attention to the Generalized System of Preferences (GSP), the per-dollar influence of multinational affiliate sales on preferential market access declines by roughly a third in the full sample, but increases by an order of magnitude when we exclude *de jure* GSP ineligible countries. Taking into account the smaller average FDI levels in GSP eligible countries, we find that a one standard deviation increase in MNE exports to the U.S. is associated with a nearly eighty percent (15 percentage point) increase in GSP access relative to the mean for GSP eligible countries. We find that this positive and significant relationship between U.S.-bound MNE sales and preferential treatment is remarkably robust across a variety of model specifications and robustness checks.

As with any empirical study, there are important caveats. The first is that our key explanatory variable, U.S.-bound MNE 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 estimates for the effect of MNE sales on U.S. trade preferences will be biased upward. By the same token, however, our results will underestimate the political economy motive (if not the implementation) if WTO non-discrimination guidelines under Article XXIV and the Enabling Clause are binding. If so, then absent these GATT provisions, we would expect even greater use of discretionary trade preferences to favor foreign affiliates of U.S. firms. Finally, we should emphasize that without detailed data on political lobbying (specifically) for trade preferences, we cannot say whether the inferred causal relationship between U.S.-bound MNE sales and trade preferences is explicitly political or simply an (apolitical) recognition of shifting national economic best interests. While it would be fascinating to decompose our findings into politically induced rent shifting versus maximization of utilitarian social welfare, such a separation is not necessary for the overall results or interpretation of our paper.

Even amidst the caveats, we conclude that discretionary trade preferences are influenced in part by the importing country's multinational firms' foreign direct investment decisions. Using a unique data set merging U.S. FDI and trade preference information, 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. GSP preferences, which are in principle designed to offer developing countries 'aid through trade' market access, seem to be particularly sensitive to MNE activities. Whether such trade policy sensitivity is good or bad depends on how the trade-investment nexus manifests itself 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.

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Appendix A. Tables

Table 1. Variables in Data Set

Variation	Variable	Source
	Dependent Variable:	
Country-Industry-Year	Preferential Access Indicator:	
	by Eligibility	4
	Effective	4
	Explanatory Variables:	
Country	Distance from U.S.	7
Country	Communist Country (yes=1)	2
Country	Terrorist Country (yes=1)	2
Country-Year	GDP	1
Country-Year	Population	1
Industry-Year	U.S. Sales	5
Industry-Year	U.S. Employment	5
Industry-Year	U.S. Import Penetration Ratio	4, 5
Industry-Year	U.S. Payroll	5
Industry-Year	Total U.S. Imports	4
Industry-Year	U.S. MFN Tariffs	4
Country-Industry-Year	Investment Data	6
Country-Industry-Year	Exports to the U.S.	3

Sources: 1. Penn World Tables; 2. CIA World Fact Book; 3. World Trade Analyzer; 4. U.S. International Trade Commission, U.S. Trade Representative; 5. U.S. Census Bureau; 6. U.S. BEA Direct Investment Abroad Database; 7. www.timeanddate.com/worldclock/distance.html; 8. U.N. Dataweb

Table 2. Data Summary Statistics: Full Sample

Variable	Mean	SD	Min	Max
Distance to U.S. (km)	8,305	3,681	734	16,357
Terrorist	.013	.115	0	1
Communist	.021	.143	0	1
Population (b.)	.051	.168	.00002	1.314
Per Capita GDP (USD)	13,547	11,776	288.4	87,825
Textile	.058	.234	0	1
Agriculture	.035	.184	0	1
U.S. Employees (m.)	.318	.550	.015	3.47
U.S. Payroll (mUSD)	8.69	7.06	.678	35.3
U.S. Sales (mUSD)	62.3	54.1	5.03	546.8
U.S. Total Imports (bUSD)	17.2	23.2	.090	214.7
MFN ad-valorem eqv. (wt)	.027	.039	0	1.36
MFN> .01	.585	.493	0	1
U.S. Import Penetration	.990	.017	.600	1.00
U.S. Number Establishments	41883	192109	92	1150793
Log Change U.S. Employees	−.031	.061	−.494	.145
Log Change U.S. Import Pen.	.001	.005	−.020	.176
C-i-t Exports to U.S. (bUSD)	.161	1.18	2.5×10^{-7}	59.2
Any GSP (indicator)	.297	.457	0	1
Any Pref (indicator)	.448	.497	0	1
Country curr. GSP el.	.573	.495	0	1
Country *de jure* GSP inel.	.388	.487	0	1
Industry curr. GSP el.	.996	.064	0	1
Any Pref Share	.198	.340	0	1
GSP Share	.122	.275	0	1
Non-GSP Share	.076	.227	0	1
RTA Share	.070	.220	0	1
El GSP (hwt)	.206	.367	0	1
El Any Pref (hwt)	.288	.392	0	1
El GSP (unwt)	.274	.333	0	1
El Any Pref (unwt)	.359	.326	0	1
MNE Sales to U.S. (bUSD)	.027	.515	0	*D*
MNE Sales to U.S. prorated by U.S. own % (bUSD)	.026	.509	0	*D*
MNE Local Sales (bUSD)	.106	.800	0	*D*
Any MNE	.209	.407	0	1
Rest of World MNE Sales (bUSD)	1.80	5.36	0	*D*

68130 Observations; 'D' Denotes BEA data redacted for confidentiality

TABLE 3. Data Summary Statistics: Reduced (De Jure GSP Eligible) Sample

Variable	Mean	SD	Min	Max
Distance to U.S. (km)	$8,297$	$3,698$	2301	$16,357$
Terrorist	.021	.144	0	1
Communist	0	0	0	0
Population (b.)	.046	.136	.00002	1.11
Per Capita GDP (USD)	$6,527$	$4,661$	288.4	$29,015$
Textile	.066	.249	0	1
Agriculture	.041	.197	0	1
U.S. Employees (m.)	.343	.585	.015	3.47
U.S. Payroll (mUSD)	8.99	7.15	.678	35.3
U.S. Sales (mUSD)	63.5	52.8	5.03	546.8
U.S. Total Imports (bUSD)	17.8	22.8	.090	214.7
MFN ad-valorem eqv. (wt)	.029	.043	0	1.36
MFN$>.01$	.570	.495	0	1
U.S. Import Penetration	.990	.015	.600	1.00
U.S. Number Establishments	47663	205273	92	1150793
Log Change U.S. Employees	$-.032$	.061	$-.494$	.145
Log Change U.S. Import Pen.	.001	.004	$-.020$	.176
C-i-t Exports to U.S. (bUSD)	.047	.400	$2.5\text{x}10^{-7}$	26.8
Any GSP (indicator)	.470	.499	0	1
Any Pref (indicator)	.547	.498	0	1
Country curr. GSP el.	.910	.287	0	1
Industry curr. GSP el.	.997	.057	0	1
Any Pref Share	.276	.377	0	1
GSP Share	.193	.326	0	1
Non-GSP Share	.083	.238	0	1
RTA Share	.078	.231	0	1
El GSP (hwt)	.327	.416	0	1
El Any Pref (hwt)	.348	.422	0	1
El GSP (unwt)	.434	.326	0	1
El Any Pref (unwt)	.461	.317	0	1
MNE Sales to U.S. (bUSD)	.0063	.1271	0	*D*
MNE Sales to U.S. prorated by U.S. own % (bUSD)	.0062	.1268	0	*D*
MNE Local Sales (bUSD)	.023	.175	0	*D*
Any MNE	.191	.393	0	1
Rest of World MNE Sales (bUSD)	2.41	5.72	0	*D*

42849 Observations; 'D' Denotes BEA data redacted for confidentiality

Table 4. Countries and Industries in Sample

NAICS-4 industries	1110, 1120, 2111, 2121, 2122, 2123, 3111, 3112, 3113, 3114, 3115, 3116, 3117, 3118, 3119, 3121, 3122, 3130, 3140, 3150, 3160, 3210, 3221, 3222, 3231, 3241, 3251, 3252, 3253, 3254, 3255, 3256, 3259, 3261, 3262, 3271, 3272, 3273, 3274, 3279, 3311, 3312, 3313, 3314, 3315, 3321, 3322, 3323, 3324, 3325, 3326, 3327, 3329, 3331, 3332, 3333, 3334, 3335, 3336, 3339, 3341, 3342, 3343, 3344, 3345, 3346, 3351, 3352, 3353, 3359, 3361, 3362, 3363, 3364, 3365, 3366, 3369, 3370, 3391, 3399
de jure GSP eligible countries	Afghanistan, Albania, Algeria, Angola, Antigua Barbuda, Argentina, Armenia, Azerbaijan, Bahrain, Bangladesh, Barbados, Belarus, Belize, Benin, Bhutan, Bolivia, Bosnia-Hercegovina, Botswana, Brazil, Bulgaria, Burkina Faso, Burundi, Cambodia, Cameroon, Cape Verde, Central African Rep., Chad, Chile, Colombia, Comoros, Congo (DROC), Congo (ROC), Costa Rica, Cote d'Ivoire, Croatia, Djibouti, Dominica, Dominican Rep., Ecuador, Egypt, El Salvador, Equatorial Guinea, Eritrea, Ethiopia, Fed. St. Micronesia, Fiji, Gabon, Gambia, Georgia, Ghana, Grenada, Guatemala, Guinea, Guinea-Bissau, Guyana, Haiti, Honduras, India, Indonesia, Iran, Iraq, Jamaica, Jordan, Kazakhstan, Kenya, Kiribati, Kyrgystan, Lebanon, Lesotho, Liberia, Libya, Macedonia, Madagascar, Malawi, Malaysia, Maldives, Mali, Mauritania, Mauritius, Moldova, Mongolia, Morocco, Mozambique, Namibia, Nepal, Nicaragua, Niger, Nigeria, Oman, Pakistan, Palau, Panama, Papua New Guinea, Paraguay, Peru, Philippines, Romania, Russia, Rwanda, Samoa, Sao Tome & Principe, Senegal, Serbia/Montenegro, Seychelles, Sierra Leone, Solomon Islands, Somalia, South Africa, Sri Lanka, St Kitts-Nevis, St Lucia, St Vincent & Grenadines, Sudan, Suriname, Swaziland, Syria, Tajikistan, Tanzania, Thailand, Togo, Tonga, Trinidad & Tobago, Tunisia, Turkey, Turkmenistan, Uganda, Ukraine, Uruguay, Uzbekistan, Vanuatu, Venezuela, Yemen, Zambia, Zimbabwe
de jure non GSP eligible countries	Australia, Austria, Bahamas, Bahrain, Belgium, Bermuda, Brunei, Canada, China, Cuba, Cyprus, Denmark, Finland, France, Germany, Greece, Hong Kong, Iceland, Ireland, Israel, Italy, Japan, Kuwait, Laos, Luxembourg, Netherlands, New Zealand, Norway, Macao, Mexico, Portugal, Qatar, Saudi Arabia, Singapore, Spain, Sweden, Switzerland, Taiwan, United Arab Emirates, United Kingdom, Vietnam
countries in both categories (due to change in status)	Czech Rep., Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovak Rep., Slovenia, South Korea

Table 5. Panel IV Results: All Preference Programs

Dependent Variable: Share of U.S. Imports Claiming Any Program Exemption

Any Pref Share	(1)	(2)	(3)	(4)
MNE sales to U.S.	.035***	.034***	.037***	.039**
[billions USD]	(.008)	(.008)	(.008)	(.017)
U.S. domestic sales	−1.19***	−1.19***	−1.19***	−1.19***
[billions USD]	(.242)	(.242)	(.242)	(.289)
total c-i-t exports to U.S.	−.002	−.002	−.003	−.003
[billions USD]	(.002)	(.002)	(.002)	(.005)
U.S. total imports (all countries)	.0004**	.0004**	.0004**	.0004*
[billions USD]	(.0002)	(.0002)	(.0002)	(.0003)
ROW MNE sales to U.S.	.0003	.0003	.0003	.0003
[billions USD]	(.0003)	(.0003)	(.0003)	(.0004)
U.S. MFN $\times$ dummy tariff>1%	4.30***	4.30***	4.29***	4.30***
[interaction term]	(0.49)	(0.49)	(0.49)	(0.71)
country GDP per capita	−.0001	−.0001	−.0001	−.0002
[in 1996 thousands of $]	(.0004)	(.0004)	(.0004)	(.0006)
country population	−.287***	−.288***	−.285***	−.285*
[in billions]	(.083)	(.083)	(.083)	(.150)
ln change in U.S. employees	.100***	.100***	.100***	.100***
	(.022)	(.022)	(.022)	(.018)
U.S. import penetration	−.282*	−.282*	−.281*	−.282*
	(.165)	(.165)	(.165)	(.154)
observations	68130	68130	68130	68130
model	IV	2SLS	GMM	GMM
instr. local sales	yes	yes	yes	yes
instr. local sales squared	no	yes	yes	yes
first-stage F-stat	27.45	17.59	17.59	3.97
Hansen's J-stat p-value	-	.334	.334	.526
cluster variable	-	-	-	ci-pair
number of clusters	-	-	-	6501

Heteroscedasticity-robust standard errors in parentheses. ***, **, * denote significance at 1%, 5%, and 10% levels of significance, respectively. Data are demeaned to remove industry- and country-fixed effects. Additional explanatory variables are the average MFN tariff, the number of employees, size of payroll, U.S. sales squared, number of establishments, and log change in the import penetration in the U.S. industry as well as year-fixed effects.

Table 6. Panel IV Results: GSP Preferences

Dependent Variable: Share of U.S. Imports Claiming GSP Tariff Exemption

GSP Share	(1)	(2)	(3)	(4)
MNE sales to U.S.	.022***	.021***	.021***	.017
[billions USD]	(.005)	(.005)	(.005)	(.011)
U.S. domestic sales	−.529***	−.529***	−.550***	−.569**
[billions USD]	(.189)	(.189)	(.189)	(.235)
total c-i-t exports to U.S.	−.003**	−.003**	−.002*	−.001
[billions USD]	(.001)	(.001)	(.001)	(.003)
U.S. total imports (all countries)	−.0003**	−.0003**	−.0003**	−.0003*
[billions USD]	(.0001)	(.0001)	(.0001)	(.0002)
ROW MNE sales to U.S.	.0004	.0004	.0003	.0003
[billions USD]	(.0003)	(.0003)	(.0003)	(.0003)
U.S. MFN × dummy tariff>1%	2.59***	2.59***	2.59***	2.61***
[interaction term]	(0.39)	(0.39)	(0.39)	(0.53)
country GDP per capita	−.001***	−.001***	−.001***	−.001**
[in 1996 thousands of $]	(.0003)	(.0003)	(.0003)	(.0005)
country population	−.308***	−.309***	−.294***	−.306**
[billions]	(.071)	(.071)	(.071)	(.124)
ln change in U.S. employees	.048***	.048***	.048***	.049***
	(.018)	(.018)	(.018)	(.015)
U.S. import penetration	−.163	−.163	−.163	−.161
	(.132)	(.132)	(.132)	(.125)
observations	68130	68130	68130	68130
model	IV	2SLS	GMM	GMM
instr. local sales	yes	yes	yes	yes
instr. local sales squared	no	yes	yes	yes
first-stage F-stat	27.45	17.59	17.59	3.97
Hansen's J-stat p-value	-	.000	.000	.025
cluster variable	-	-	-	ci-pair
number of clusters	-	-	-	6501

Heteroscedasticity-robust standard errors in parentheses. ***, **, * denote significance at 1%, 5%, and 10% levels of significance, respectively. Data are demeaned to remove industry- and country-fixed effects. Additional explanatory variables are the average MFN tariff, the number of employees, size of payroll, U.S. sales squared, number of establishments, and log change in the import penetration in the U.S. industry as well as year-fixed effects.

Table 7. Panel IV Results: Excluding *de jure* GSP Ineligible Countries

Dependent Variable: Share of U.S. Imports Claiming Exemption Under:

	GSP Only		Any Preference Program	
	(1)	(2)	(3)	(4)
MNE sales to U.S.	.513***	.515***	.442***	.444***
[billions USD]	(.137)	(.130)	(.139)	(.129)
U.S. domestic sales	−.866***	−.864***	−1.76***	−1.76***
[billions USD]	(.285)	(.284)	(.341)	(.339)
total c-i-t exports to U.S.	−.116***	−.117***	−.086***	−.086***
[billions USD]	(.030)	(.026)	(.032)	(.026)
U.S. total imports (all countries)	−.0004*	−.0004*	.0007**	.0007**
[billions USD]	(.0002)	(.0002)	(.0003)	(.0003)
ROW MNE sales to U.S.	.0006	.0006	.0008	.0008
[billions USD]	(.0005)	(.0005)	(.0005)	(.0005)
U.S. MFN × dummy tariff>1%	7.46***	7.46***	9.06***	9.06***
[interaction term]	(.66)	(.66)	(.79)	(.79)
country GDP per capita	−.006***	−.006***	−.002	−.002
[in 1996 thousands of $]	(.001)	(.001)	(.002)	(.002)
country population	.920***	.921***	.909***	.909***
[billions]	(.099)	(.099)	(.109)	(.108)
ln change in U.S. employees	.062**	.062**	.121***	.121***
	(.028)	(.028)	(.033)	(.033)
U.S. import penetration	−.443*	−.441*	−.813***	−.813***
	(.245)	(.245)	(.271)	(.270)
observations	42849	42849	42849	42849
model	2SLS	GMM	2SLS	GMM
instr. local sales	yes	yes	yes	yes
instr. local sales squared	yes	yes	yes	yes
first-stage F-stat	11.69	11.69	11.69	11.69
Hansen's J-stat p-value	.948	.948	.978	.978

Heteroscedasticity-robust standard errors in parentheses. ***, **, * denote significance at 1%, 5%, and 10% levels of significance, respectively. Data are demeaned to remove industry- and country-fixed effects. Additional explanatory variables are the average MFN tariff, the number of employees, size of payroll, U.S. sales squared, number of establishments, and log change in the import penetration in the U.S. industry as well as year-fixed effects.

Table 8. IV Tobit Results

	Full Sample		Reduced Sample	
	Any Pref Share	GSP Share	Any Pref Share	GSP Share
	(1)	(2)	(3)	(4)
MNE sales to U.S.	.066***	.364***	.287	.588**
[billions USD]	(.016)	(.102)	(.276)	(.270)
U.S. domestic sales	.283	.660**	.236	.720**
[billions USD]	(.215)	(.299)	(.276)	(.298)
total c-i-t exports to U.S.	.006	−.057**	−.008	−.099*
[millions USD]	(.005)	(.023)	(.054)	(.052)
U.S. total imports (all countries)	.002***	−.001	.002***	−.001*
[billions USD]	(.000)	(.001)	(.001)	(.001)
ROW MNE sales to U.S.	.002***	.002*	.003**	.002
[billions USD]	(.001)	(.001)	(.001)	(.001)
U.S. MFN tariff rate	2.34***	1.75***	2.69***	1.76***
[trade weighted ad-valorem equivalent]	(.08)	(.10)	(.10)	(.10)
country-fixed effects	yes	yes	yes	yes
industry-fixed effects	yes	yes	yes	yes
year-fixed effects	yes	yes	yes	yes
instr. local sales	yes	yes	yes	yes
observations	68130	68130	42849	42849
left censored obs. (dep. var.= 0)	37612	47924	19410	22724
right censored obs. (dep. var.= 1)	3122	1959	2997	1953

Standard errors in parentheses.***, **, * denote significance at 1%, 5% ,and 10% levels of significance, respectively.

Table 9. 3SLS Results

	Full Sample		Reduced Sample	
	Any Pref Share	GSP Share	Any Pref Share	GSP Share
FDI Equation				
preference measure	.586***	1.79***	.019	.223***
	(.146)	(.36)	(.027)	(.050)
local sales	.141***	.140***	.170***	.156***
[billions USD]	(.004)	(.005)	(.006)	(.007)
local sales squared	.002***	.002***	-.029***	-.027***
	(.0001)	(.0002)	(.001)	(.001)
total c-i-t exports to U.S.	.196***	.198***	.176***	.180***
[billions USD]	(.002)	(.002)	(.001)	(.002)
country population	-.530***	-.172	-.181***	-.380**
[billions]	(.135)	(.201)	(.045)	(.062)
Preference Equation				
MNE sales to U.S.	.036***	.022***	.343***	.426***
[billions USD]	(.008)	(.007)	(.098)	(.086)
U.S. domestic sales	.138*	.162***	.021	.0002
[billions USD]	(.081)	(.050)	(.127)	(.090)
total c-i-t exports to U.S.	-.003	-.004**	-.061***	-.089***
[billions USD]	(.002)	(.002)	(.018)	(.016)
U.S. total imports (all countries)	-.00005	-.0005***	.0004*	-.0009***
[billions USD]	(.0002)	(.0001)	(.0002)	(.0001)
ROW MNE sales	-.004***	-.005***	-.001*	-.002***
[billions USD]	(.0004)	(.0003)	(.0007)	(.0005)
U.S. MFN $\times$ dummy tariff>1%	1.67***	.98**	4.88***	4.69***
[interaction term]	(.53)	(.45)	(.80)	(.70)
country population	-.293***	-.314***	.928***	.916**
[billions]	(.085)	(.072)	(.104)	(.091)
ln change in U.S. employees	.144***	.062***	.188***	.087***
	(0.023)	(0.016)	(0.035)	(0.025)
Observations	68130	68130	42849	42849
$\mathcal{X}_1^2$	32706.7	18297.3	19239.4	14677.3
$\mathcal{X}_2^2$	2828.1	1902.4	2204.1	1399.4

Standard errors in parentheses; * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 10. Alternative Specifications: Full Sample

Reported: Coefficient Estimates for MNE Sales

Model	Any Pref Share (1)	GSP Share (2)	El Any (3)	El GSP (4)	Non-GSP Share (5)
Panel IV overidentified GMM	$.04^{***}$	$.02^{***}$	$.02^{**}$	.004	$.01^{**}$
IV Tobit with c-, i-, y- fixed effects	$.07^{***}$	$.36^{***}$	.01	$.24^{**}$	$-.001$
3SLS	$.04^{***}$	$.02^{***}$	$.02^{*}$	.01	$.01^{***}$
Pooled IV [unclustered/2-way clustering]	$.04^{***/ns}$	$.02^{***/**}$	$.02^{*/ns}$	$-.01^{**/ns}$	$.02^{ns/ns}$
Panel OLS	$.01^{***}$	$.002^{*}$	$-.004^{**}$	$-.0004$	$.01^{***}$
Pooled OLS [unclustered/2-way clustering]	$.02^{***/***}$	$.01^{***/**}$	$.005^{**/ns}$	$.003^{ns/ns}$	$.01^{***/**}$
Panel IV: joint country-industry FE overidentified GMM	$.005^{*}$	$.005^{**}$	.0003	$.009^{**}$	.0003
Panel IV: joint industry-year FE overidentified GMM	$.03^{***}$	$.02^{***}$	$.02^{***}$	$.06^{*}$	$.01^{***}$
Panel IV: joint country-year FE overidentified GMM	$.04^{***}$	$.02^{***}$	$.02^{***}$	$-.0003$	$.02^{***}$

'ns' denotes results not statistically significant from zero at the 10% level. ***, **, * denote significance at 1%, 5%, and 10% levels, respectively. Where noted for pooled runs, significance is reported for both robust and two-way (country and industry) cluster-robust standard errors, respectively.

Table 11. Alternative Specifications: Reduced Sample

Reported: Coefficient Estimates for MNE Sales

Model	Any Pref Share (1)	GSP Share (2)	El Any (3)	El GSP (4)	Non-GSP Share (5)
Panel IV overidentified GMM	.44***	.52***	.46***	.51***	-.07
IV Tobit with c-, i-, y- fixed effects	.29	.59**	.21	.33	.65*
3SLS	.34***	.43***	.37***	.41***	−.08
Pooled IV [unclustered/2-way clustered]	.76***/**	1.22***/**	1.35***/**	1.35***/**	−.45***/**
Panel OLS	.11***	.10***	.08***	.10***	.01
Pooled OLS [unclustered/2-way clustering]	.08***/ns	.05***/**	.04ns/ns	.04ns/ns	.02ns/ns
Panel IV: joint country-industry FE overidentified GMM	.24**	.07	.30**	.44***	.15***
Panel IV: joint industry-year FE overidentified GMM	.24***	.26***	.27***	.29*	−.02
Panel IV: joint country-year FE overidentified GMM	.30***	.29***	.31***	.32***	.008

'ns' denotes results not statistically significant from zero at the 10% level. ***, **, * denote significance at 1%, 5%, and 10% levels, respectively. Where noted for pooled runs, significance is reported for unclustered robust and two-way (country and industry) cluster-robust standard errors, respectively.