U.S. MULTINATIONALS AND PREFERENTIAL MARKET ACCESS

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PRELIMINARY

Abstract. This paper examines 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 create a three-way panel spanning 80 industries, 184 countries, and ten years (1997-2006). Consistent with existing theory, we find that off-shoring multinational activity and preferential market access are positively and consistently correlated, both in the pooled sample and within countries, industries, and years. Using instrumental variables to account for the 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 4 to 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 developing 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 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 vertical or 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 have set tariffs cooperatively since the inception of the General Agreement on Tariffs and

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

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 off-shoring 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 (PTAs) 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 extent that multinationals operate horizontal 'tariff jumping' operations abroad, those activities will have either a negligible or small *positive* effect on the investment-source

[2]While in principle Article XXIV and GSP limit the extent of unilateral discretion, the supposed uniformity (across industries and/or countries) required of MFN exemptions appears to be quite weak in practice. (This is easily confirmed by a cursory examination of the U.S. tariff code.) Indeed, were FTAs truly non-discriminatory across industries (as stipulated by Article XXIV) – or were GSP preferences in practice determined only by country and industry (but without discretion at the level of the country-industry pair) – then our results would not withstand country and industry fixed effects, as they do.

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 (U.S. foreign subsidiaries' goods sales to the U.S.) apart from horizontal import-substituting investment (subsidiaries' sales to the 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 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 whether export oriented

[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.

FDI by U.S. firms causes 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 exactly 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 4-5 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

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

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. Finally, 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. Since they consider a small country framework in which world and domestic prices for given MFN tariff are fixed, the government will grant preferential market access if and only if foreign lobby contributions more than offset the welfare losses from foregone tariff revenue. Summarizing, while each of these earlier studies captures important elements of preferential tariff setting, none consider the potential influence of international investment.

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, N. 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 number of foreign trading partners; α_i denotes the welfare weight on 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:

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

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 $k\times1$ 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 our pooled sample, 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 (equivalent to the 6-digit NAICS level), the investment data are available only at the

[8]Note that the error term will capture 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.

more aggregated 4-digit NAICS level. For this 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 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 the same qualitative 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%. In the panel specification, we interact the linear MFN term with the dummy 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. As we specified in (4.3), however, we have potentially three different types of fixed effects. We control for time-fixed effects by introducing year indicator variables, whereas we remove industry- and country-fixed effects γ_i and γ_c by demeaning (this also removes any variables that depend only on the country or the industry).[11] Yet once we have demeaned, we can no longer estimate a Tobit model, but rather have to assume a linear probability model, thus ignoring possible mass points for θ_{cit} at 0 and 1.[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 sales to the *local market*.[13] In a just-identified specification of the model, local MNE

[11]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]In a robustness check for the linear panel model, we censor the data at 0 and find our results to be quantitatively consistent (see Table 7). Very few observations (roughly 5%) are censored at 1 to begin with.

[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 rules out 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

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

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 (6-digit NAICS) 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 weighted[16] and straight (unweighted) averages across the relevant 6-digit NAICS 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.).

be independent of U.S. trade preferences. (On average, less than 25% of firms in our sample sell both locally and to the US.) Indeed, in a simple pooled OLS version of the model, we found that export oriented MNE sales were positive 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.

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

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 (and virtually all changes under GSP) are effective July 1st, 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. (Both CNL implementation and product-country specific CNL waivers 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.) Finally, highly restrictive rules of origin restrictions 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 of several preferential codes by program, industry, and country of origin.[17] We use this information to construct our baseline measure of the preferential treatment variable 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 a preferential program. Note that the effective

[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.

preference measure offers an additional feature in that it does not require an ad-hoc weighting scheme to aggregate from 6-digit to 4-digit NAICS.[18]

Our foreign investment data are from the U.S. Bureau of Economic Analysis international accounts for U.S. direct investment abroad. The BEA-DIA 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.[19] 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 (equivalent to 6-digit HTS), so that each 4-digit NAICS industry includes multiple 8-digit HTS product lines.[20] 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.[21] 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

[18]It should be noted that some of the deviation between the USITC and USTR 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 robustness of our results to either preference variable definition suggests that this caveat is of minor consequence in the context of our study.

[19]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 US MNE sales, and virtually all sales when pro-rated by percentage of US ownership.

[20]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.

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

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-DIA data.

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 US, 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[22] 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.[23] 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

[22]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.

[23]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.

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.

6. Results

Table 3 reports the estimation results when the dependent variable is the share of imports by industry i from country c in year t for which any preferential tariff treatment was claimed. Country and industry specific fixed effects are removed by demeaning,[24] and time fixed-effects are incorporated as year dummies. The table reports coefficient estimates for only the variables of major interest. Explanatory variables included but not reported are exporting country population, 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.[25] 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.[26] In particular, higher overall imports and lower country-industry-year exports to the U.S. 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 MFN average tariff rate is positive and also significant at the 1% level; recall that the theoretical prediction

[24] Standard errors are corrected accordingly.

[25] 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%); as one would expect, negligible MFN rates have little effect on preferences, while higher MFN tariffs seem to be important.

[26] Of the other included variables, the following are statistically significant at the 10% level: country population (billions) [$-.29$ ($p < .001$)], U.S. establishments (thousands) [$.0002$ ($p = .02$)], U.S. employees (millions) [$-.08$ ($p = .02$)], the MFN tariff [-3.04 ($p < .001$)] and U.S. sales squared (bUSD) [2.4 ($p < .001$)].

is ambiguous. U.S. MNE sales to the rest of the world sales appear to have no effect on the rate of preferential market access, with a p-value of .6. 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. 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 4.3 percentage points. Conditional on having any FDI initially,[27] a one standard deviation increase in the mean level of export oriented MNE sales for a given country and industry then implies more than a four percentage point jump in the rate of preferential imports – roughly a twenty 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 19.35.[28] 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, indicating that the instruments appear to be valid. 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 and hovering around a four and a half percentage point increase in the rate of preferential market access for each $1 billion in MNE sales to the U.S.

In Table 4, 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.

[27]Among countries with any U.S. FDI, the mean FDI sales to the U.S. (prorated by U.S. ownership) is .094 billion USD with a standard deviation of .97 billion USD.

[28]The first stage coefficient for the instrument (MNE local sales) is .176 ($p < .001$).

We see that the effect of U.S. multinational sales continues to be statistically significant, but the estimate is somewhat smaller: an increase in multinational sales to the U.S. of $1 billion leads to a statistically significant (at the 1% level) increase in the exemption share of roughly three percentage points. Again, the coefficient estimates are robust across each specification of the model. The implication seems to be that while GSP preferences appear to be subject to offshoring MNE activity, the per-dollar effect of FDI is roughly two thirds that for 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.[29] Also of concern is the much higher value of Hansen's J test for instrument validity – it seems that for the GSP variable in the full sample, the instrument may be correlated with the second stage error term.

In Table 5, we further restrict the sample to exclude *de jure* GSP ineligible countries under the 1974 Trade Act.[30] In the reduced sample, 91% of observations are for currently GSP eligible countries, opposed to 57% in the full sample. The effect of removing the *de jure* GSP ineligible countries is striking: the coefficient estimates increase by approximately an order of magnitude across the board. The instruments are somewhat

[29]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 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.

[30]Countries are immediately deemed ineligible if they are members of the EU, 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.

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. In interpreting the coefficient estimates, it is important to keep in mind, too, that the average level of MNE sales to the U.S. is much lower in the reduced sample. Conditional on having any FDI,[31] a one standard deviation increase in MNE sales to the U.S. would induce roughly a 17 percentage point expansion of the rate of GSP access – a ninety percent increase relative to the (reduced) sample average rate of GSP preferential access of .19.

We now turn in Table 6 to an IV Tobit specification of the model for the full sample, 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. Our IV Tobit model therefore is pooled across countries and industries, though we continue to include time dummies. With a pooled sample, we can now include time-invariant country and industry control variables, but we also face the potential complication of error term correlation within country or industry groups. We thus correct our standard errors for two-way clustering at the country and industry levels following the technique for non-linear models specified by Cameron, Gelbach, and Miller (2006). Finally, due to computational limitations in the maximum likelihood estimation of the clustered IV Tobit model and the redundancy among several of our earlier control variables for U.S. industry characteristics, we reduce the set of controls to include only those explicitly specified by the model, the key gravity regressors – distance, country population, and country income per capita – and a dummy variable to indicate *de jure* GSP ineligible countries.

The first two columns of Table 6 consider the dependent variable measured as preference program shares from the Customs data, as in the earlier tables, while the second two columns consider preference program eligibility (aggregated from 6- to 4-digit NAICS

[31]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 .038 billion USD with a standard deviation of .27 billion USD.

using historical (1997) trade flows as weights). The last column returns to the measurement of preferential market access by share of imports entering under a preferential tariff code, this time restricting attention to only those imports that enter under non-GSP preference programs.

The IV Tobit coefficient estimates are generally higher in absolute value than for commensurate runs under the linear panel IV model. All of the five presented specifications give the predicted positive sign for the FDI variable, but the coefficient is statistically significant at the 10% level for only the baseline measure, percentage of imports entering under any preference program. (If we cluster by only one dimension – by industry *or* by country – or if we neglect to cluster at all, then the coefficients are statistically significant across the board.) The coefficient signs and magnitudes are generally consistent with the linear version of the panel model, though again the statistical significance is lower due to multi-dimensional clustering. Finally, note that most of the censoring is at zero (55% of the observations for the 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 performs quite well in the full sample (less so in the reduced sample). The key finding, of course, is that U.S. export oriented activity abroad seems to increase the rate of preferential market access, controlling for the endogeneity of export oriented sales and country, industry, and year fixed effects. Under the linear panel IV specification, a $1 billion increase in MNE sales to the U.S. is associated with roughly a four 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 increase in MNE sales to the U.S. is nearly four times as great, in terms of the percentage point increase in the rate of preferential market access, as the effect of a one standard deviation increase in MNE sales for the average country in the full sample.

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. Conditional on there being any preference accorded, the effect of $1 billion in MNE sales back to the U.S. (a one standard deviation increase conditional on having any U.S. MNE activity) is a remarkable 6 to 22 percentage points, the upper bound of which would double the rate of preferential access relative to the sample mean of .20. At the same time, however, the statistical significance of the key regressor is substantially less in the IV Tobit specification (with two-way clustering); as a consequence, we are cautious not to overstate the findings from this non-linear specification of the model.

7. Alternative Specifications

We conduct a battery of robustness checks. In a first step, we consider a variety of alternative specifications of the model. As argued earlier, we find the most sensible starting point to be the instrumented versions of the model presented in the previous section: either the linear panel specification, or the pooled version under Tobit.[32] For comparison, however, we also ran the IV version of the model on the pooled data (without country and industry fixed effects) and for the reduced form (no IV) linear versions of the model under pooled and panel OLS. All pooled estimations include year fixed effects and are corrected for two way clustering at the country and industry levels. (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.) Finally, as a back

[32]The ideal, of course, would be to include all industry and country fixed effects in the IV Tobit model, but this procedure proves computationally infeasible given the dimensionality of the data.

of the envelope check on the IV Tobit results in which we could account for country and industry fixed effects, we created a linear panel IV version of the model with data censored for observations for which any measure of preference program eligibility was strictly greater than zero.

We report the results for each of these alternative models in Tables 7 (for the full sample) and 8 (for the reduced sample), for each of the alternative definitions of the dependent variable. We also include the already reported baseline results from the linear panel IV and IV Tobit specifications for ease of reference. For brevity, the tables list only the coefficient of interest, MNE sales to the U.S., for each specification. 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.

Reviewing the tables, the coefficient estimates are largely consistent with the findings reported previously. Notice that the instrumental variables results are very similar across specifications (linear panel IV, linear pooled IV, and censored linear panel IV – the IV Tobit results are a bit higher), as are the no-IV findings (panel and pooled OLS). Comparing columns 1 and 3 for all preference programs, and columns 2 and 4 for GSP, we see that the estimates are similarly consistent across different 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 measures in columns 3 and 4).[33] Finally, note that while the coefficient estimates vary, they are consistently positive and significant at the

[33]As an additional robustness check, we confirm that the trade weights are not responsible for our findings in columns (3) and (4). In the baseline panel IV specification, we find that the key coefficient estimates for *unweighted* versions of the dependent variables El Any and El GSP are virtually identical: for El Any, the key coefficient estimate for the trade-weighted version is .0268**, while that for the unweighted version is .0276***; for El GSP the key estimate for the trade weighted version of the dependent variable is .005, compared to .006 for the unweighted version. Removing the trade weights has a slightly larger effect in the reduced sample: for the panel IV specification in Table 8, the key coefficient estimates in columns (3) and (4) fall from .57*** to .47*** and from .64** to .51*** respectively; the qualitative results, however, remain unchanged.

1 percent level for all but a few variants of the model. Comparing results across the two tables, we again suggest focusing on the full sample results for the estimates of all preference programs in general, and on the reduced sample results for GSP programs specifically.

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 6-digit NAICS sub-aggregates (in contrast to the trade weighted version used in the benchmark model), and we measure program eligibility using *unweighted* aggregation from 6- to 4-digit NAICS. For each of these variants, our baseline results remain qualitatively unchanged, though of course the standard errors increase substantially for some of the smaller sample sizes. In a last, perhaps foolhardy investigation, we introduce country-industry specific fixed effects into our linear IV model, and thus rely solely on intertemporal variation as a means of identification. Remarkably, we still find that the coefficient of our key explanatory variable of interest, MNE sales to the U.S., is positive and significant at the 10% level, though the coefficient estimate falls by roughly two-thirds, to .01 with a p-value of .06.[34] The bottom line is simple – we put our model and results through the wringer, and find the estimates remarkably robust: offshoring-type export oriented investment by U.S. multinationals appears to induce an economically and statistically significant increase in preferential market access to the U.S. for investment host countries and industries.

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

[34]Results from the baseline over-identified GMM panel IV specification.

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 4 to 5 percentage point increase in the rate of preferential duty free access from all preferential programs combined. A one-standard deviation increase in MNE export oriented sales to the US would induce roughly a twenty percent increase in market access relative to the mean. Restricting attention to 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 US is associated with a nearly ninety percent increase in GSP access (17 percentage points) relative to the mean for GSP eligible countries. The positive and significant relationship between FDI and preferential treatment is remarkably robust across a variety of model specifications and robustness checks.

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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

Variable	Mean	SD	Min	Max
Distance to U.S. (km)	8,305	3,681	734	16,357
Terrorist	.013	.11	0	1
Communist	.021	.14	0	1
Population (b.)	.051	.17	.00002	1.314
Per Capita GDP (USD)	13,546	11,775	288.4	87,825
Textile	.058	.23	0	1
Agriculture	.035	.18	0	1
U.S. Employees (m.)	.318	.55	.015	3.47
U.S. Payroll (mUSD)	8.69	7.06	.678	35.32
U.S. Sales (mUSD)	62.3	54.10	5.032	546.81
U.S. Total Imports (bUSD)	17.2	23.24	.090	214.74
MFN ad-valorem eqv. (wt)	.027	.04	0	1.357
MFN ad-valorem eqv.	.032	.04	0	.757
MFN$>.01$	.585	.49	0	1
U.S. Import Penetration	.990	.02	.599	.999
Log Change U.S. Employees	.020	.09	1.119	.450
Log Change U.S. Import Pen.	.001	.005	.019	.176
C-i-t Exports to U.S. (bUSD)	.161	1.18	2x10^{-7}	59.17
Any GSP	.300	.46	0	1
Any Pref.	.448	.50	0	1
Country curr. GSP el.	.573	.49	0	1
Country *de jure* GSP inel.	.388	.49	0	1
Industry curr. GSP el.	.996	.06	0	1
All Pref Share	.198	.34	0	1
GSP Share	.122	.26	0	1
Non-GSP Share	.076	.23	0	1
RTA Share	.070	.22	0	1
Elig. GSP (hwt)	.206	.37	0	1
Elig. Any Pref (hwt)	.288	.39	0	1
Elig. GSP	.274	.33	0	1
Elig. Any Pref	.359	.33	0	1
MNE Sales to U.S. (bUSD)	.027	.52	0	*D*
MNE Sales to U.S. prorated by U.S. own % (bUSD)	.020	.44	0	*D*
MNE Local Sales (bUSD)	.106	.80	0	*D*
Any MNE	.209	.41	0	1
Rest of World MNE Sales (bUSD)	1.79	5.36	0	63.423

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

Table 3. Panel IV Results: All Preference Programs

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

	(1)	(2)	(3)	(4)
MNE sales to U.S.	$.043^{***}$	$.043^{***}$	$.044^{***}$	$.045^{**}$
[billions USD]	(.011)	(.011)	(.010)	(.021)
U.S. domestic sales	-1.24^{***}	-1.24^{***}	-1.24^{***}	-1.24^{***}
[billions USD]	(.244)	(.244)	(.244)	(.292)
total cit exports to U.S.	$-.003$	$-.003$	$-.003$	$-.003$
[billions USD]	(.003)	(.003)	(.002)	(.005)
U.S. total imports (all countries)	$.0005^{**}$	$.0005^{**}$	$.0005^{**}$	$.0005^{**}$
[billions USD]	(.0002)	(.0002)	(.0002)	(.0003)
ROW MNE sales to U.S.	.0004	.0004	.0004	.0004
[billions USD]	(0.0003)	(0.0003)	(0.0003)	(0.0003)
U.S. MFN $\times$ dummy tariff>1%	4.29^{***}	4.29^{***}	4.29^{***}	4.29^{***}
[interaction term]	(0.50)	(0.50)	(0.50)	(0.71)
country GDP per capita	$-.0001$	$-.0001$	$-.0001$	$-.0001$
[in 1996 thousands of $]	(.0004)	(.0004)	(.0004)	(.0004)
ln change in U.S. employees	$.101^{***}$	$.101^{***}$	$.100^{***}$	$.101^{***}$
	(.022)	(.022)	(.022)	(.018)
U.S. import penetration	$-.290^{*}$	$-.290^{*}$	$-.290^{*}$	$-.290^{*}$
	(.165)	(.165)	(.165)	(.155)
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	19.35	14.76	14.76	3.94
Hansen's J-stat p-value	-	.722	.722	.828
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 export country's population, the number of employees, size of payroll, US sales squared, number of establishments, and log change in the import penetration in the U.S. industry as well as time-fixed effects.

TABLE 4. Panel IV Results: GSP Preferences

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

	(1)	(2)	(3)	(4)
MNE sales to U.S.	.027***	.027***	.032***	.026*
[billions USD]	(.007)	(.007)	(.007)	(.014)
U.S. domestic sales	−.550***	−.550***	−.557***	−.581***
[billions USD]	(.191)	(.191)	(.190)	(.237)
total cit exports to U.S.	−.004**	−.004**	−.004**	−.003
[billions USD]	(.002)	(.002)	(.002)	(.003)
U.S. total imports (all countries)	−.0002	−.0002	−.0002	−.0002
[billions USD]	(.0001)	(.0001)	(.0001)	(.0002)
ROW MNE sales to U.S.	0.0002	0.0002	0.0002	0.0002
[billions USD]	(0.0002)	(0.0002)	(0.0002)	(0.0003)
U.S. MFN × dummy tariff>1%	2.58***	2.58***	2.59***	2.60***
[interaction term]	(0.39)	(0.39)	(0.39)	(0.54)
country GDP per capita	−.0011***	−.0011***	−.0011***	−.0012***
[in 1996 thousands of $]	(.0003)	(.0003)	(.0003)	(.0005)
ln change in U.S. employees	.049***	.049***	.048***	.049***
	(.018)	(.018)	(.018)	(.015)
U.S. import penetration	−.167	−.167	−.165	−.164
	(.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	19.35	14.76	14.76	3.94
Hansen's J-stat p-value	-	.001	.001	.065
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 export country's population, the number of employees, size of payroll, US sales squared, number of establishments, and log change in the import penetration in the U.S. industry as well as time-fixed effects.

Table 5. 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.	.626***	.637***	.536***	.545***
[billions USD]	(.202)	(.192)	(.198)	(.182)
U.S. domestic sales	−.948***	−.943***	−1.90***	−1.90***
[billions USD]	(.291)	(.289)	(.348)	(.346)
total cit exports to U.S.	−.112***	−.115***	.081**	.084**
[billions USD]	(.035)	(.029)	(.035)	(.027)
U.S. total imports (all countries)	.0000	.0000	.0011***	.0011***
[billions USD]	(.0003)	(.0003)	(.0003)	(.0003)
ROW MNE sales to U.S.	.0002	.0002	.0009*	.0009*
[billions USD]	(.0004)	(.0004)	(.0005)	(.0005)
U.S. MFN × dummy tariff>1%	7.43***	7.43***	9.12***	9.12***
[interaction term]	(0.67)	(0.67)	(0.80)	(0.80)
country GDP per capita	−.005***	−.005***	−.001	−.001
[in 1996 thousands of $]	(.001)	(.001)	(.002)	(.002)
ln change in U.S. employees	.068**	.069**	.127***	.127***
	(.030)	(.030)	(.034)	(.034)
U.S. import penetration	−.449*	−.446*	−.829***	−.827***
	(.247)	(.247)	(.273)	(.273)
observations	41697	41697	41697	41697
model	2SLS	GMM	2SLS	GMM
instr. local sales	yes	yes	yes	yes
instr. local sales squared	yes	yes	yes	yes
first-stage F-stat	7.14	7.14	7.14	7.14
Hansen's J-stat p-value	.861	.861	.903	.903

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 export country's population, the number of employees, size of payroll, US sales squared, number of establishments, and log change in the import penetration in the U.S. industry as well as time-fixed effects.

Table 6. IV Tobit Results

	All Pref Share	GSP Share	El Any	El GSP	Non-GSP Share
	(1)	(2)	(3)	(4)	(5)
MNE sales to US	.223*	.190	.097	.073	.058
[billions USD]	(.116)	(.141)	(.070)	(.171)	(.064)
U.S. domestic sales	.262	−.494	.147	.302	.968*
[billions USD]	(.787)	(.942)	(1.22)	(1.16)	(.564)
total cit exports to US	.005	−.089***	.017	−.047	.029*
[millions USD]	(.013)	(.034)	(.018)	(.055)	(.016)
U.S. total imports (all countries)	−.004	−.005*	−.002	−.005	.000
[billions USD]	(.002)	(.003)	(.003)	(.004)	(.002)
ROW MNE sales to US	−.002	.001	.008	.006	−.004
[billions USD]	(.005)	(.006)	(.006)	(.007)	(.004)
U.S. MFN tariff rate	2.87***	2.31***	3.19**	3.32***	2.45***
[trade weighted ad-valorem equivalent]	(.89)	(.84)	(1.27)	(1.26)	(.58)
country GDP per capita	−.005	−.015	.002	−.019	−.001
[in 1996 thousands of $]	(.006)	(.011)	(.004)	(.015)	(.006)
country population	.218	.257	.262	.351	.157
[billions]	(.166)	(.190)	(.160)	(.246)	(.113)
distance from U.S.	−.030***	.010	−.019***	−.024*	−.077***
[thousands km]	(.010)	(.009)	(.007)	(.012)	(.015)
observations	68130	68130	68130	68130	68130

Standard errors clustered by both country and industry in parentheses. ***, **, * denote significance at 1%, 5% ,and 10% levels of significance, respectively. The excluded instrument is local sales. Time dummies and a dummy for *de jure* GSP program ineligibility are included, but not reported.

Table 7. Alternative Specifications: Full Sample

Reported: Coefficient Estimates for MNE Sales

Model	All Pref Share (1)	GSP Share (2)	El Any (3)	El GSP (4)	Non-GSP Share (5)
Panel IV overidentified GMM	$.04^{***}$	$.03^{***}$	$.03^{**}$	$.01$	$.01^{**}$
Pooled IV [unclustered/2-way clustering]	$.04^{***/-}$	$.03^{***/-}$	$.03^{**/-}$	$-.02^{***/ns}$	$.02^{*/ns}$
Censored Panel IV: overidentified GMM	$.04^{***}$	$.03^{***}$	$.03^{***}$	$.004$	$.01^{**}$
IV Tobit [unclustered/2-way clustering]	$.22^{***/*}$	$.19^{***/-}$	$.10^{***/-}$	$.07^{-/ns}$	$.06^{***/ns}$
Panel OLS	$.01^{***}$	ns	$-.004^{***}$	ns	$.01^{***}$
Pooled OLS [unclustered/2-way clustering]	$.02^{***/***}$	$.004^{***/*}$	$.004^{*/ns}$	ns	$.01^{***/***}$

'ns' denotes results not statistically significant from zero at the 20% 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 8. Alternative Specifications: Reduced Sample

Reported: Coefficient Estimates for MNE Sales

Model	All Pref Share	GSP Share	El Any	El GSP	Non-GSP Share
	(1)	(2)	(3)	(4)	(5)
Panel IV overidentified GMM	.55***	.64***	.57***	.64**	ns
Pooled IV [unclustered/2-way clustered]	.91***/−	1.49***/*	1.65***/−	1.66***/−	−.58***/ns
Cen. Pan IV: overidentified GMM	.73**	.89***	.66**	.76***	ns
IV Tobit [unclustered]	3.18***	4.66***	4.90***	5.43***	1.27**
Panel OLS	.12***	.09***	.07***	.09***	ns
Pooled OLS [unclustered/2-way clustering]	.07**/−	.04**/**	ns	ns	.03*/ns

Estimates reported only if p value is less than .2; 'ns' denotes results not statistically significant from zero at the 20% 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.