Optimal Tariffs: The Evidence*

Christian Broda
University of Chicago, GSB and NBER

Nuno Limão
University of Maryland
NBER and CEPR

David E. Weinstein
Columbia University and NBER

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Abstract

The theoretical debate over whether countries can and should set tariffs in response to the foreign export elasticities they face goes back to Edgeworth (1894). Despite the centrality of the optimal tariff argument in trade policy, there exists no evidence about whether countries actually exploit their market power in trade by setting higher tariffs on goods that are supplied inelastically. We estimate disaggregate foreign export supply elasticities and find evidence that countries that are not members of the World Trade Organization systematically set higher tariffs on goods that are supplied inelastically. The typical country in our sample sets tariffs 9 percentage points higher in goods with high market power relative to those with low market power. This large effect is of a magnitude similar to the average tariffs in the data and market power explains more of the tariff variation than a commonly used political economy variable. The result is robust to the inclusion of other determinants of tariffs and a variety of model specifications. We also find that U.S. trade restrictions that are not covered by the WTO are significantly higher in goods where the U.S. has more market power. In short, we find strong evidence that these importers have market power and use it in setting non-cooperative trade policy.

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* Contact information: cbroda@chicagogsb.edu, Limao@wam.umd.edu, Dew35@columbia.edu. Broda and Weinstein would like to thank the NSF for generous funding under grant NSF #0214378. Limão gratefully acknowledges the excellent research assistance of Piyush Chandra and the financial support of the IMF research department where he was a resident scholar during part of this research. We wish to thank Stephanie Aaronson, Fernando Alvarez, Kyle Bagwell, Alan Deardorff, Peter Debaere, Bill Ethier, John Romalis, Robert Staiger two anonymous referees and the editor for extremely useful and detailed comments. We also thank seminar participants at various institutions for numerous comments and suggestions (CEPR trade meeting Summer 2006, Chicago Fed, Empirical Investigations in Trade 2006, Dartmouth College, Harvard University, International Monetary Fund, Midwest International Economics Spring 2006, NBER ITI meeting Winter 2006, Princeton University, Johns Hopkins, Syracuse University, University of Chicago, University of Virginia, University of Wisconsin). The views expressed in this paper are those of the authors.

Introduction

The idea that countries set tariffs in response to their market power in international markets is the single most controversial result in international trade policy. It is not hard to find examples of first class theorists arguing that it provides the underlying motive for the world trading system (Bagwell and Staiger, 1999) while others argue that it is little more than an intellectual curiosity with no practical value in all but the largest countries (Krugman and Obstfeld, 1997). Given that the theoretical debate over optimal tariffs goes back over a century, one might ask, "What evidence is there in favor or against the notion that tariffs vary inversely with export supply elasticities?" The answer is none.

In this paper, we provide evidence that non-members of the World Trade Organization (WTO) systematically set higher import tariffs on goods in which they have market power, i.e. goods that are supplied inelastically. We also find that U.S. trade restrictions that are not covered by the WTO are significantly higher in goods where the U.S. has more market power. The results are robust to the inclusion of political economy variables and a variety of model specifications. The results isolate an effect that is not only statistically significant but also economically important both relative to other explanations and to the average tariff in the typical country. In short, we find strong evidence that countries have market power in imports and exploit it in setting their trade policy.

The theory that a country might gain from protection has a long history.[1] The intuition for why countries might gain from tariffs through an improvement in their terms-of-trade stems from two key insights. The first, from Torrens (1833) and Mill (1844), is that there are many possible prices at which countries would be willing to trade. The imposition of a tariff creates a welfare loss due to consumption and production distortions, but it can also produce a gain if foreign suppliers reduce their prices in order to maintain market access. If the losses due to the domestic distortion are less than the gains from the price or terms-of-trade effect, a country can gain from a tariff.

Edgeworth (1894) provided the key insight regarding when a country should impose a tariff. He showed that as long as a foreign country's offer curve was not perfectly elastic, a country could gain from a tariff. In this case, the reduction in import demand caused by a tariff leads to a reduction in the price of all units imported and this first order gain offsets the distortion losses from lower imports. Bickerdike (1907) extended Edgeworth's framework and developed the formula relating the welfare maximizing tariff and the inverse of the export supply elasticity. Although Bickerdike framed his derivation with one import good and a welfare maximizing government, the basic insight that a country's "optimal" tariff is increasing in its market power applies to more general settings and does

[1] Irwin (1996) carefully examines the history of thought on protection and the next two paragraphs draw on it .

not require governments to maximize welfare, as we discuss in the theory section. Our objective in this paper is to quantify how important market power is in determining a country's tariff structure. We also refer to this effect as the terms-of-trade motive for tariffs, as is common in the literature.

Trade economists have long been uncomfortable with the optimal tariff argument. From a normative perspective, the key objection is that if a tariff improves a country's terms-of-trade, it worsens those of its trading partner, who may therefore retaliate leaving both worse off relative to free trade.[2] As a positive theory of trade protection, the optimal tariff argument is often questioned for two reasons. First, "small countries have very little ability to affect the world prices of either their imports or other exports, so that the terms-of-trade argument is of little practical importance" as Krugman and Obstfeld (1997, p. 226) write in their undergraduate textbook. But until now, there is little evidence on this account, and it may be more correct to argue, as Feenstra (2004) does in his textbook, that the basic welfare maximizing formula for tariffs "is not very helpful because there is little that we know empirically about the elasticity of foreign export supply." The second objection is that governments do not set tariffs to maximize social welfare. While this last argument is often true, we have already noted that it is not a necessary condition for a positive relationship between tariffs and market power.

Despite the skepticism regarding the practical importance of the optimal tariff argument, it continues to feature prominently in the leading theoretical trade policy models. Grossman and Helpman (1995) extend their endogenous trade policy model to the case where a country is "large", i.e. it faces finite export supply elasticities. Although not stressed in their paper, there would be no motive for trade talks in their model in the absence of a terms-of-trade use of the tariff. This is a key point made by Bagwell and Staiger (1999) who provide an economic theory of the General Agreement on Trade and Tariffs (GATT). In this and in subsequent work, Bagwell and Staiger have strongly argued that the use of tariffs to explore the terms-of-trade effect can explain many of the key features of the current multilateral trading system. Their work has been quite influential despite the fact that there was no direct evidence that countries used, or indeed possessed, market power in trade prior to entering into reciprocal liberalization in the GATT or its successor, the WTO. In fact, Rose (2004) uses 68 measures of aggregate trade policy and finds no significant effect of liberalization upon GATT/WTO accession. Since our initial working paper however, independent research by Bagwell and Staiger

[2] This outcome is stressed by Scitovsky (1942) but Johnson (1953-54) shows that certain countries may actually gain from using optimal tariffs even with retaliation.

(2006) provides evidence for their theory by showing that WTO accession leads to greater tariff reductions in products with higher initial import volumes.[3]

There is some evidence that changes in trade policy affect the prices of the goods that countries import.[4] This evidence generally attributes the effect to imperfect competition in specific industries. More importantly, these studies do not argue or estimate whether countries changed their trade policies to affect their terms-of-trade much less if they did so taking the export supply elasticity into account. In fact, little is known empirically about the foreign export supply elasticity since most calculations of trade elasticities simply assume that it is infinite.

The measurement of foreign export supply elasticities, which quantifies an importers' implied market power, constitutes one of the contributions of this paper. We rely on the methodology of Feenstra (1994) and Broda and Weinstein (2006) to estimate these elasticities for each 4-digit Harmonized System (HS) category during the period 1994-2003 for the 15 non-WTO members for which this and the relevant tariff data is available for a large fraction of products.

We find that the inverse export supply elasticity faced by an importer is between 1 and 3 for the typical 4-digit HS good. We also test several conjectures about these elasticities and find support for them in our estimates. For example, larger countries face less elastic export supply curves, which indicates that, on average, they have more market power than small ones. This is true if we use GDP or a country's share in world imports in a particular good as a measure of size. Moreover, these elasticities are positively correlated across importing countries for any given good. This is likely to be the case if importers systematically have more market power for some types of goods. We confirm this conjecture by finding that importers face much flatter export supply curves for commodities, where the inverse elasticity is 0.5, than for differentiated products, where it is 2.4. The implied pass-through rates from tariffs to export prices are also in line with existing evidence on tariff and exchange rate pass-through.

Using these elasticities we then estimate that, prior to entering the WTO, countries set higher tariffs on products where they have more market power, i.e. higher inverse export supply elasticities. This effect is present both when we compare median tariff rates across countries and when we compare actual tariff rates across Harmonized Tariff System (HS) 4-digit goods within countries and industries. The impact of market power on tariffs is robust to many different specifications. The effect is present

[3] The optimal tariff equilibrium is also used as the threat point and the main theoretical motive for preferential trade agreements in many influential papers.

[4] Kreinin (1961) estimates that more than two-thirds of U.S. tariff reductions in the Geneva trade Round were passed on as higher prices to countries exporting to the U.S. Chang and Winters (2002) estimate that the elimination of internal tariffs between Argentina and Brazil caused prices of exports into Brazil to fall. There is also evidence of imperfect pass-through from exchange rates and that the effect is symmetric to that of tariff changes for the U.S. auto sector (Feenstra, 1989).

using continuous and discontinuous versions of the export supply elasticity measure and controlling for unobserved industry heterogeneity in *each* country. The estimate is positive and significant in the pooled sample and also positive in all 15 countries and significant for 13. Moreover, we address the possibility of omitted variable bias and measurement error via an instrumental variables approach.

The result is also robust to the inclusion of variables that capture two prominent motives for protection: revenue and lobbying. As is common in recent tests of political economy models (e.g., Goldberg and Maggi, 1999), we find that the lobbying effect is strong. Nonetheless, the market power effect on tariffs remains positive and significant. It is at least as important as the lobbying motive both in terms of the magnitude and the fraction of tariff variation explained.

The estimated effect is also economically important. In particular, we find that the typical country sets tariffs 9 percentage points higher in goods where it has medium or high market power relative to those with low market power. These goods represent two-thirds of each country's sample. The effect is important in 13 of the 15 non-WTO countries; in China it is 35 percentage points. This is roughly the same magnitude of China's average tariff over all goods and the same relationship between the effect and the average tariff holds for the typical country. We estimate that removing this motive for tariff setting would lead to significant increases in the prices received by foreign exporters, particularly those selling in the larger countries in our sample: China, Russia and Taiwan.

In order to follow the theory closely, we focus on countries that are not members of the WTO and thus set their policies in a unilateral, non-cooperative way. However, we also analyze the role of market power in shaping a subset of trade policies that are determined non-cooperatively by the U.S., a large member of the WTO. The U.S. sets non-tariff barriers and statutory tariffs (i.e. rates it applies to some non-WTO members) with few or no restrictions from the GATT/WTO. We find that market power is also an important determinant of these trade policies the U.S. sets unilaterally. Interestingly, we find no such effect on those U.S. tariffs set according to WTO rules. This finding is broadly consistent with Bagwell and Staiger's theory of the GATT/WTO and it suggests that market power would play an important role for *all* U.S. trade policies if they were set non-cooperatively, e.g. in the absence of the WTO. More generally, these results show that the importance of the terms-of-trade motive extends beyond non-WTO members and so understanding its impact on trade policy is essential.

The paper is organized as follows. We first present the basic theory that we test. In section 3, we describe the estimation methodology for the elasticities. In section 4, we describe the data and assess the validity of the elasticity estimates. We present the estimation results for non-WTO members in section 5 and for the U.S. in section 6. We conclude in section 7.

2. Theory

The basic theory underlying the optimal tariff argument is well established. Therefore, in this section, we provide the basic intuition for the result and show how it is robust to the inclusion of political economy considerations. We are interested in how a country sets policy in the absence of agreements. So we focus on a country that takes as given the policies of the remaining $n \geq 1$ countries.

Suppose each individual has a utility defined over a numeraire good, c_0, and a vector of non-numeraire goods $u(\mathbf{c})$. Here we consider the simpler case where $u(\mathbf{c})$ is separable and in the appendix we consider an alternative that matches our estimation approach. Omitting the country subscript, we write this individual's utility as

$$U = c_0^h + \sum_g u_g(c_g^h). \tag{1}$$

Each individual h with income I^h chooses expenditure on each good c_g to maximize (1) subject to $c_0^h + \sum_g p_g c_g^h \leq I^h$, where p_g is the domestic price for c_g. Given the quasilinearity and separability, the demand for each good g is simply a function of its own price, i.e. $c_g = c_g(p_g)$. An individual's indirect utility can be written as $v^h = I^h + \sum_g \psi_g(p_g)$, where the last term represents consumer surplus.[5] Social welfare is then the sum of the individual indirect utilities:

$$W = \sum_h \left[I^h + \sum_g \psi_g\left(p_g\right) \right]. \tag{2}$$

To determine income, I^h, we employ the standard assumptions in the leading endogenous trade policy models, e.g. Grossman and Helpman (1994, 1995). First, the numeraire is freely traded and produced using only labor according to a constant returns production. So the equilibrium wage is determined by the marginal product in this sector, which we normalize to one. Second, the non-numeraire goods are produced using a constant returns production with labor and one factor specific to the good. So each specific factor earns a quasi-rent that is increasing in the good's price, $\pi_g(p_g)$. Finally, tariff revenues for each good, $r_g(p_g)$, are redistributed uniformly to all individuals. All individuals own a unit of labor and a fraction of them also own up to one unit of specific capital. So, if we normalize the population to one, the wage income also equals one and we can rewrite social welfare as

$$W = 1 + \sum_g \left[\pi_g\left(p_g\right) + r_g\left(p_g\right) + \psi_g\left(p_g\right) \right]. \tag{3}$$

The world price for each traded good $g \in G_m$ is determined by the market clearing conditions

$$m_g\left(\left(1+\tau_g\right)p_g^*\right) = m_g^*\left(p_g^*\right) \quad \forall g \in G_m \tag{4}$$

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[5] More specifically, $\sum_g \psi_g(p_g) \equiv \sum_g \left[u_g\left(c_g\left(p_g\right)\right) - p_g c_g\left(p_g\right) \right]$.

where m_g represents home's import demand written as a function of the domestic price, $p_g = (1+\tau_g)p_g^*$, and m_g^* is the rest of the world's export supply. From this we obtain prices as functions of the trade policy: $p_g\left(\tau_g\right)$, $p_g^*\left(\tau_g\right)$.[6]

A government choosing the optimal tariff to maximize (3) will then set the tariff for each good g according to the following first order condition:[7]

(5) $$\tau_g p_g^* \frac{dm_g}{d\tau_g} - m_g \frac{dp_g^*}{d\tau_g} = 0 \quad \forall g \in G_m.$$

The first term represents the domestic distortion caused by the tariff. The second term represents the terms-of-trade effect. If the country has no market power in trade, i.e. if the export supply elasticity is infinite, then $dp_g^* / d\tau_g = 0$, and the optimal tariff is zero. Otherwise the optimal tariff is positive and can be shown to equal the inverse export supply elasticity,[8]

(6) $$\tau_g^{opt} = \omega_g \equiv \left[\left(dm_g^* / dp_g^*\right)\left(p_g^* / m_g^*\right)\right]^{-1}.$$

The positive relationship between protection and the inverse elasticity, ω_g, extends to more general settings, one of which we examine in the appendix. Here we highlight a few other points. The separability assumption in our model implies that the tariffs in (6) do not reflect any monopoly power in the export sector. The original Bickerdike formulation allowed for both market power in the import and export sectors. He showed in a two good world that if a country could not impose an export tax, the optimal tariff was linearly related to the inverse export elasticity (although not with a coefficient of unity) plus another term that was related to the inverse demand elasticity of this country's exports. Graaf (1949-50) extended this result to multiple goods and showed that if countries can impose both export taxes and import tariffs, and the cross-price elasticities are all zero, then the optimal policy is to impose an export tax equal to the inverse demand elasticity and import tariff equal to the inverse export elasticity. The bottom line from these more complex policy experiments is that monopoly power in the export sectors may create an additional motive for the use of import tariffs (c.f. Alvarez and Lucas, 2005; Gros, 1987 shows this is the case even for "small" countries when products are

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[6] In a setting with many importers the equilibrium prices also depend on other importers' tariffs. This does not affect the results here because the optimal tariff prediction takes the other countries' policies as given *and* we will focus on the case where there is a constant foreign export supply elasticity, which is independent of prices.

[7] Taking $dW/d\tau_g$=0; using the envelope theorem, $\frac{d\psi_g(p_g)}{d\tau_g} = -c_g \frac{dp_g}{d\tau_g}$; and $\frac{dp_g}{d\tau_g} = \left(1+\tau_g\right)\frac{dp_g^*}{d\tau_g} + p_g^*$, we obtain (5).

[8] By applying the implicit function theorem to (4) we obtain an expression for $dp_g^* / d\tau_g$, which can be used in (5) to obtain the expression in (6) after some algebraic manipulation.

differentiated). But this additional motive does not eliminate the first order incentive to impose higher tariffs in sectors in which imports are supplied less elastically. It is this last prediction that we test.

A common objection to the terms-of-trade motive for tariffs as a positive theory of trade policy is that governments do not choose tariffs to maximize social welfare. However, a positive relationship such as the one in equation (6) can also describe the equilibrium policy even in models where the government has other objectives, such as redistribution of income to particular specific factor owners. A key insight in the trade policy literature is that of targeting (e.g. Bhagwati and Ramaswami, 1963), which states that if a government has domestic objectives they can be met more efficiently using a instruments other than tariffs, and that when these instruments target the distortion at the source, the optimal tariff is zero in a small economy. The counterpart to this insight for a large economy is that when the government's objective function places some value on additional income from the improved terms-of-trade, we obtain a positive relationship between the tariff rate and the inverse elasticity provided that the government also uses instruments such as subsidies or transfers that target the other externalities.[9]

The positive relationship between tariffs and inverse elasticities can also hold if the government's objective is not social welfare maximization *and* it does not have other policies to redistribute income to producers. Grossman and Helpman (1995) extend their political contributions trade model to the large country case. The non-cooperative tariff that the government chooses in that model maximizes a weighted sum of social welfare and contributions, C_g, from the L organized lobbies representing specific factor owners, i.e. $aW+\sum_{g\in L}C_g$. In this case, the optimal tariff is

$$\tau_g^{GH} = \omega_g + \frac{I_g - \alpha}{a+\alpha}\frac{z_g}{\sigma_g} \tag{7}$$

where the last term reflects the lobbying motive for tariffs. If a, the government's marginal rate of substitution between contributions and social welfare, is infinite then we obtain the welfare maximizing optimal tariff. More importantly, the partial positive relationship between the tariff and ω_g holds even when this government places *no* weight on social welfare. We can see this by noting that as long as any positive fraction of the population is organized into lobbies, i.e. $\alpha \in (0,1]$, the second term in (7) remains finite even if the government does not value social welfare at all, i.e. if a equals zero. This occurs because lobbies' contributions account for all the costs and benefits of the set

[9] Most trade policy models that provide a political economy motive for a tariff must in fact rule out these other instruments; otherwise in those models the tariff's only role would be to affect the terms-of-trade. This is an example of the puzzle of the use of inefficient policies for redistribution. Rodrik (1995) points out it is particularly problematic for trade policy, Drazen and Limão (Forthcoming) provide one explanation for it.

of tariffs they bid on. One such benefit is the terms-of-trade gain that the lobbies reap via the redistributed tariff revenue, even if they produce a good other than *g*.

The tariff for an organized group is increasing in z_g, the inverse import penetration ratio, because a given tariff generates larger benefits for a factor owner if it applies to more units sold.[10] The tariff depends negatively on the import demand elasticity, σ_g, reflecting the basic Ramsey taxation intuition that, once the terms-of-trade effect is accounted for, the tariff's distortion is increasing in this elasticity. As Helpman (1997) shows, the size and elasticity effect captured by z_g/σ_g also arises in other political economy models and so we will use this variable as one of the controls in the estimation.

The key obstacle in estimating the impact of market power on tariffs is obtaining elasticity estimates for a broad set of countries and goods. In order to achieve this, we must impose some structure on the data. We now briefly describe how the standard approach above can be extended in a way that is both compatible with our estimation of the elasticities *and* delivers the positive effect of market power on tariffs.

In the next section, we describe the system of import demand and export supply equations that we use to estimate the elasticities. This system can be derived in a setting where any foreign variety (i.e. a good imported from a particular exporter) is valued according to a CES utility function, and supply is perfectly competitive. In the appendix, we show that the optimal tariff in a model with CES utility over foreign varieties of a given good is identical to equation (6), i.e. the inverse export elasticity. This occurs when utility is separable across goods (but not varieties); the tariffs do not affect the relative demand of varieties within any given good; and hence the only distortion that is addressed by the tariff is the terms-of-trade externality. As we prove in the appendix, there are three assumptions that imply the tariff in a good does not affect the relative demand of varieties within it; these assumptions are mainly driven by the constraints imposed by the data, sample and estimation. First, consumption and foreign export supply elasticities within any given good are constant. Second, they are identical across varieties, i.e. exporters of that good. Third, tariffs of a given country in any given year are equal across exporters of the same good.[11]

[10] The variable z_g is defined as the ratio of domestic production value to import value, where the latter excludes tariffs.

[11] In an extreme version of the optimal tariff argument, we may expect countries to discriminate across different exporters of the same good. The most common way this occurs is through preferential agreements. However, in the country-year sample we consider such agreements are not important. Nine of the fifteen countries do not report preferential rates and for five countries that do report them, those rates apply to only a small share of their tariff lines or imports. One reason why these countries generally define tariffs on the basis of a good and not origin is that the administrative cost of optimal discrimination may be too high since it would require preventing international arbitrage by strictly enforcing rules of origin.

3. Estimating Foreign Export Supply and Import Demand Elasticities

A key reason why the impact of market power on tariffs has not been examined before is the difficulty of obtaining reliable measures of the elasticity of foreign export supply as required by equations (6) and (7) for example. In fact, most calculations of trade elasticities simply assume that countries face an infinitely elastic supply of exports and therefore estimate only import demand elasticities. In this section, we explain how to obtain the elasticities of foreign export supply and import demand for each good in each importing country. We do so using a methodology derived by Feenstra (1994) and extended by Broda and Weinstein (2006).[12]

We estimate the import demand and *inverse* export supply elasticities (σ_{ig} and ω_{ig} , respectively), using the following system of import and export equations:

$$(8) \qquad \Delta^{k_{ig}} \ln s_{igvt} = -\left(\sigma_{ig} - 1\right)\Delta^{k_{ig}} \ln p_{igvt} + \varepsilon_{igvt}^{k_{ig}}$$

$$(9) \qquad \Delta^{k_{ig}} \ln p_{igvt} = \frac{\omega_{ig}}{1+\omega_{ig}} \Delta^{k_{ig}} \ln s_{igvt} + \delta_{igvt}^{k_{ig}}$$

Equation (8) represents the optimal demand of country *i* for a given variety *v* of a good *g* – derived from a CES utility function – and (9) represents the residual export supply country *i* faces in that variety. Both are expressed in terms of shares, where s_{igvt} is the share of variety *v* of good *g* in country *i*. The equation for each variety imported by country *i* is differenced with respect to time *t* and a benchmark variety of the same good *g* imported by *i*, denoted k_{ig}. More specifically, the difference operator we use for the shares and domestic prices is defined as $\Delta^{k_{ig}} x_{igvt} = \Delta x_{igvt} - \Delta x_{igk_{ig}t}$ where Δ stands for a simple time difference. The last parameter in (8) , $\varepsilon_{igvt}^{k_{ig}} = \varepsilon_{igvt} - \varepsilon_{igk_{ig}t}$, represents demand shocks that differ across varieties, for example ε_{igvt} includes *changes* in taste or quality for a variety *v*. Similarly, $\delta_{igvt}^{k_{ig}} = \delta_{igvt} - \delta_{igk_{ig}t}$, where δ_{igvt} includes shocks to the residual export supply when expressed as a function of importer prices, e.g. exchange rate changes between countries *i* and *v*.

An important feature of the method used is that it requires double log-differencing the data. This implies that the *level* of tariffs on varieties or goods will not affect our estimated elasticities, which reduces the possibility of reverse causality when we estimate their effect on tariffs. More generally, *changes* in trade costs at the good level will also typically not affect the estimates. Similarly,

[12] Broda and Weinstein (2006) estimate import demand elasticities for a range of imports but do not report the export supply elasticities. Feenstra (1994) reports both elasticities for eight specific products. Both studies focus only on the U.S. Iriwin (1988) and Romalis (Forthcoming) report both elasticities. However, because they are at the aggregate level and for only two countries (the U.K. and U.S. respectively), they cannot be used to estimate the impact of market power on tariffs.

our estimates will not be biased in the presence of any fixed costs of exporting that cause some set of countries to have no exports of a variety over time whereas others have positive exports. Finally, our estimates will be unbiased even if there are quality differentials across countries for a given variety or if there are good specific trends in these differentials.[13]

There are two important conditions needed to identify the elasticities. First, ω_{ig} and σ_{ig} are constant over varieties and this time period (but they can vary over importers and goods). Second, demand and supply shocks *relative* to the benchmark variety are uncorrelated, i.e. $E_t\left(\varepsilon_{igvt}^{k_{ig}}\delta_{igvt}^{k_{ig}}\right)=0$.

In the robustness section and in the appendix, we analyze if our results are sensitive to some of these identifying assumptions; for now we simply note why they may be plausible. The elasticity of substitution over varieties of a good, σ_{ig}, is a preference parameter and thus not likely to vary across the short time period we examine or across varieties for a finely defined good. The residual export supply elasticity, ω_{ig}, depends, among other things, on production elasticities and on the rest of the world's import demand elasticities, $\sigma_{j\neq ig}$. The latter should not change much over the time-span of our data, 6-9 years, for the reason noted above. However, we will test whether allowing for different elasticities across exporters of a given good changes the results. Finally, the assumption of independence of relative errors is likely to be reasonable because the large shocks on a yearly frequency are often due to bilateral exchange rate changes. These are captured as supply shocks in (9) and, at this frequency, they are unlikely to be correlated with demand shocks such as relative taste or quality. Ultimately, this is an empirical question and in the appendix we test and find evidence that supports this assumption.

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[13] Generally the foreign export supply is written as a function of the price received by the exporter, p_{igvt}^{*}. In the presence of some ad valorem trading cost, τ_{igvt}, and bilateral exchange rate, e_{ivt}, the domestic price is $p_{igvt}=\left(1+\tau_{igvt}\right)e_{ivt}\,p_{igvt}^{*}$ and so $\Delta^{k_{ig}}\ln p_{igvt}=\Delta^{k_{ig}}\ln p_{igvt}^{*}+\Delta^{k_{ig}}\ln e_{ivt}+[\Delta\ln\left(1+\tau_{igvt}\right)-\Delta\ln\left(1+\tau_{igk_{ig}t}\right)]$. Thus the export supply error, $\delta_{igvt}^{k_{ig}}$, contains the bilateral exchange rate shocks, as noted in the text. Since these can be frequent and large they are a more important source of variation than shocks to relative trade costs. For example, if τ_{igvt} represents transport costs then a change in it is likely to be similar across varieties of a good when it is due to say improved importer ports, so the relative shock is zero. If τ_{igvt} represents a tariff, then the *relative* shock in brackets is also zero in several cases. To see this note that we define the good at the HS4 level. Although most of these countries' MFN tariffs are set at the HS6 level only 10% of their tariff variation occurs within HS4 so often we have $\tau_{igvt}\approx\tau_{igk_{ig}t}$. The relative trade cost shock is zero even if the country is undertaking a unilateral liberalization (e.g. China), since its tariff changes are similar for all its exporters of a given good, or if it implemented a preferential agreement prior to time *t* (e.g. Saudi Arabia). Some of the countries in the sample implemented preferential agreements during the period we use the trade data, which is a relative shock that the differencing does not eliminate. However, this is reflected only on the export supply error, $\delta_{igvt}^{k_{ig}}$, since the demand equation controls for the domestic price, and thus does not invalidate our elasticity identification assumption. Nonetheless, we address the possibility that such tariff changes affect the elasticity estimate by using IV in the tariff estimation section.

To take advantage of the independence of errors condition, $E_t\left(\varepsilon_{igvt}^{k_{ig}}\delta_{igvt}^{k_{ig}}\right)=0$, we solve (8) and (9) in terms of those errors and multiply them together to obtain:

$$(10)\qquad \left(\Delta^{k_{ig}}\ln p_{igvt}\right)^2=\theta_{i1}\left(\Delta^{k_{ig}}\ln s_{igvt}\right)^2+\theta_{i2}\left(\Delta^{k_{ig}}\ln p_{igvt}\Delta^{k_{ig}}\ln s_{igvt}\right)+u_{igvt}$$

where $\theta_{ig1}=\dfrac{\omega_{ig}}{\left(1+\omega_{ig}\right)\left(\sigma_{ig}-1\right)}$, $\theta_{ig2}=\dfrac{\omega_{ig}\left(\sigma_{ig}-2\right)-1}{\left(1+\omega_{ig}\right)\left(\sigma_{ig}-1\right)}$ and $u_{igvt}=\dfrac{\varepsilon_{igvt}^{k_{ig}}\ \delta_{igvt}^{k_{ig}}}{\sigma_{ig}-1}$. Note that the error term, u_{igvt}, is correlated with the regressands that depend on prices and expenditure shares. However, Feenstra (1994) shows that a consistent estimator of $\boldsymbol{\theta}_{\mathbf{ig}}=\left(\theta_{ig1},\theta_{ig2}\right)$ can be obtained by averaging (10) over time. To see this we can write the "between" version of (10) as:

$$(11)\qquad \bar{Y}_{igv}=\theta_{ig1}\bar{X}_{1,igv}+\theta_{ig2}\bar{X}_{2,igv}+\bar{u}_{igv}$$

where $Y_{igvt}=\left(\Delta^{k_{ig}}\ln p_{igvt}\right)^2$, $X_{1,igvt}=\left(\Delta^{k_{ig}}\ln s_{igvt}\right)^2$, $X_{2,igvt}=\left(\Delta^{k_{ig}}\ln p_{igvt}\Delta^{k_{ig}}\ln s_{igvt}\right)$ and bars on top of these variables denote their time averages (the *t* subscript is dropped). The independence of errors assumption implies $E_v\left(\bar{u}_{igv}\right)=0$. Intuitively, the time-series identification problem of a single importer-good pair is solved by using the information available in all the varieties imported of that good. While data on prices and shares of a single variety can pin down a relationship between σ_{ig} and ω_{ig}, they are insufficient to determine the exact value of these elasticities. Additional varieties of the same importer-good pair provide information about how these elasticities are related, and given that the true σ_{ig} and ω_{ig} are assumed constant across varieties of the same good, this information helps estimate the elasticities.

Feenstra (1994) also notes that provided there are three varieties of the same importer-good pair that are sufficiently different in their second moments then the true underlying elasticities are exactly identified.[14] We will slightly modify this criterion and follow the procedure in Broda and Weinstein (2006). They show that in the presence of measurement error in the prices used to compute unit values for each variety, an additional term needs to be added to (10) and a different weighting scheme should be used to estimate (11). In particular, unit values are generally better measured when based on large volumes. Therefore, the weights and the additional term are inversely related to the quantity imported of the variety and the number of periods the variety had positive imports. This implies that at least four varieties per good are needed to obtain identification.

[14] The relative variance of demand to supply shocks cannot be identical across varieties. Otherwise they would describe the same relationship between the relevant elasticities and no information is added by having a second variety.

Using this weighting scheme, we first estimate (11) to obtain $\hat{\boldsymbol{\theta}}_{\mathbf{ig}}$ and check that it implies elasticities in the set of economically feasible estimates, i.e., $\sigma_{ig} > 1$ and $\omega_{ig} > 0$ for all *i* and *g*. If they don't then we perform a grid search over the feasible values of $\boldsymbol{\theta}_{\mathbf{ig}}$. We evaluate the sum of squared errors of (11) at values of $\sigma_{ig} > 1$ and $\omega_{ig} > 0$ at intervals that are approximately 5 percent apart.[15]

The precision for the typical elasticity is obtained by bootstrapping. We re-sampled the data for each importer-good pair 250 times and computed estimates of the importer-good elasticity each time. The procedure used to compute these bootstrapped elasticities replicates the one in the estimation of the actual elasticities.[16]

4. Data, Descriptive Statistics and Assessment of Elasticity Estimates

4.1 Data

In order to estimate the impact of market power we need data on tariffs, domestic production and elasticities. In deciding what set of countries to include we face both theoretical and empirical constraints. The theory applies to countries setting their trade policy unilaterally in a non-cooperative way. Since a major function of the GATT/WTO is to allow countries to reciprocally lower their tariffs in order to internalize the terms-of trade effects, we focus the test on non-GATT/WTO members. In section 6 we provide additional evidence for a set of policies set non-cooperatively by a WTO member.

Our tariff data comes from the TRAINS database, which provides data at the 6-digit HS level. Unfortunately, some non-WTO countries report this data for only a small share of goods making it impossible to make meaningful comparisons across goods or compute country averages. Therefore, we focus only on the fifteen non-WTO countries that report tariffs in at least one third of all 6-digit goods. The set of countries and the years we use are reported in Table 1.[17]

Our sample includes a non-negligible part of the world economy and is representative of the world as a whole in some dimensions. It includes countries from most continents. The average per capita GDP in the sample is \$9,000, which is similar to the world average of \$8,900. The 15 countries

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[15] For computational easiness, we performed the grid search over values of σ_{ig} and ρ_{ig} where ρ_{ig} is related to ω_{ig} in the following way: $\omega_{ig} = \rho_{ig} / \left(\sigma_{ig}\left(1-\rho_{ig}\right)-1\right)$. The objective function was evaluated at values for $\sigma_{ig} \in [1.05, 131.5]$ at intervals that are 5 percent apart, and for $\rho_{ig} \in [0.01, 1]$ at intervals 0.01 apart. Only combinations of σ_{ig} and ρ_{ig} that imply $\sigma_{ig} > 1$ and $\omega_{ig} > 0$ are used. To ensure we used a sufficiently tight grid, we cross-checked these grid-searched parameters with estimates obtained by non-linear least squares as well as those obtained through Feenstra's original methodology. Using our grid spacing, the difference between the parameters estimated using Feenstra's methodology and ours differed only by a few percent for those σ_{ig} and ω_{ig} for which we could apply Feenstra's "between" approach.

[16] The only difference is that a broader grid is used in case the regression coefficients imply elasticities of the wrong sign. This is solely for computational easiness, since this bootstrap procedure has to compute over 3 million elasticities.

[17] This criteria was binding for only four countries: Bahamas, Brunei, Seychelles and Sudan.

comprise 25% of the world's population and close to 20% of its GDP (in PPP). This is due to the fact that it includes two of the world's ten largest economies, China and Russia, as well as several smaller but non-negligible countries such as Taiwan, Ukraine, Algeria, Saudi Arabia and Czech Republic.

The trade data is obtained from the United Nations Commodity Trade Statistics Database (COMTRADE). This database provides quantity and value data at 6-digit 1992 HS classification for bilateral flows between all countries in the world. As we can see from Table 1, the import data for most countries in our sample covers the period 1994-2003. For Taiwan we use UNCTAD's TRAINS database since COMTRADE does not report data for this country.

4.2 Descriptive Statistics

The choice of what constitutes a good is dictated by data availability. The more disaggregated the choice of good the fewer varieties per good we have, and thus at some point the elasticity estimates become too imprecise. Therefore, in estimating (8) and (9) we define a good, *g*, as a 4-digit HS category and a variety, *v*, as a 6-digit good from a particular exporter. Table 2 shows that the typical country has 1100 4-digit categories with positive imports between 1994 and 2003. The typical good in the sample is imported from 17 different countries. There are between 15,000 and 66,000 varieties being imported per year by each of these countries. For instance, there were 40 different varieties of live fish (4-digit HS 0301) imported by China in 2001, among them were "trout" (HS 030191) from Australia and "eels" (HS 030192) from Thailand. The high degree of specialization of exports suggests that one should be cautious about assuming that the share of a country in world GDP is a sufficient proxy for the ability of a country to gain from a tariff. If China places a tariff on live fish, it is not clear that Thailand can easily export its eels elsewhere and receive the same price.

Table 2 also shows statistics describing the tariff data at the HS4 level. There are several important features to note. First, variation across countries accounts for one-third of the total variation. The mean across countries ranges from 4 to 38 percent, with 10 being the typical value, the range and typical values for medians are similar to the mean. Second, there is also considerable variation within countries, the standard deviation ranges from 1 to 26 percent and 9 is the typical value. Finally, since we estimate the elasticities at the HS 4 level we aggregate the tariff data up to that level by taking simple averages. As we can see from the last column the precise aggregation method and focus on HS 4 variation has little impact since over 90% of the variation in tariffs for the typical country occurs across HS 4 rather than within it.

If one were to take size, as measured by GDP, as a good proxy for market power then the data on tariff levels suggests that the skepticism regarding the optimal tariff argument is not entirely unwarranted. First, as we can see in Table 2, although China is both the largest country in our sample and has the highest tariff, Taiwan, the third largest country has a below average tariff. The correlation between median tariff and the log of GDP is 0.48 and that between average tariffs and GDP is 0.53 However, if we drop China, those correlations fall to 0.05 and 0.10, respectively.

Data on the within-country variation also suggests that the tariff setting policies are likely to be more complex than a simple application of the optimal tariff calculus. Figure 1 portrays the within-country frequency distribution of tariffs at the 4-digit level. Although most countries have large dispersion across goods, there are three with either little dispersion, such as Bolivia, or some dispersion but with most tariffs grouped into certain value bins, Oman, and Saudi Arabia. Moreover, we observe truncation and some bunching at the lower end of distribution, where about 9% of all tariffs are zero.

There are a couple of important implications of the stylized facts mentioned above. First, although considering cross-country results may yield interesting insights, it may be more reasonable to focus on the effect of market power in determining tariffs across goods within countries. Second, in some countries the data seems to militate against a simple relationship in which policymakers equate the tariff level with a continuous variable such as export elasticities or degrees of political power. One can imagine many reasons for this. Perhaps policy makers are uncertain of inverse elasticities or political connectedness and therefore divide their tariff schedule in various categories rather continuous levels; maybe policymakers employ other means of protection at their disposal when they want to achieve high levels of protection; maybe countries are averse to setting tariffs too high out of fear of retaliation; or maybe once tariffs hit prohibitive levels, there is no reason to raise them further.

All of these complications suggest that the effect of market power on tariffs may not follow the exact functional forms postulated by simple and stylized models. Thus our focus will not be to test if the data confirms or rejects the optimal tariff theory expressed in a particular functional form, but rather to estimate the impact of market power on tariffs.

4.3 Elasticity Estimates

Since we conduct the analysis at the 4-digit level for each country, we estimate over 12,000 foreign export supply elasticities – far too many to present individually. Therefore, in Table 3A we report their summary statistics. In theory, the inverse foreign export elasticity, ω_{ig}, can be anywhere between zero and infinity. So the median provides a useful way to characterize the estimates, as it is less sensitive to extreme values. The median inverse elasticity across all goods in any given country ranges from 0.9 to 3. It is 1.6 in the full sample, implying a median elasticity of supply of 0.6, that is a 1 percent increase in prices elicits a 0.6 percent increase in the volume of exports for the typical good.

As will become clear it is also useful to consider how different the typical estimates are across terciles. The table shows that the typical estimate for low market power goods (i.e. those with inverse elasticities in the bottom 33rd percentile of a given country) is 0.3, about five times smaller relative to medium market power goods (1.6) and 180 times smaller than high market power goods (54).

Obviously, some of the 12,000 elasticities are imprecisely estimated. Tables 3A and 3B indicate that this appears to be an issue for the largest estimates, as indicated by how much higher the mean is relative to the median and by the wider bootstrap confidence intervals for elasticities in the top decile. In fact, when we trim the top decile of the sample, the means fall by almost an order of magnitude, down to 13. The same is true for the standard deviation. However, the key point to keep in mind is that even in the trimmed sample there is considerable variation in market power across goods within a country that is not driven by measurement error and can be used to estimate the effect of market power on tariffs.

Since the standard errors are non-spherical we assess the precision of the estimates via bootstrapping. More specifically, we resample the data and compute new estimates for each of the elasticities 250 times.[18] Since there is no simple way to describe the dispersion of all estimates, we focus on the key question for our purpose. Namely, whether the estimates are precise enough to distinguish between categories of goods in which a country has low versus medium or high market power. If the answer to this question is positive then we can address measurement error by using this categorical variable as either our market power measure or as an instrument for the continuous variable.

Before describing the results for the full sample consider the following specific case for Russia where we divide goods into a low, medium or high market power category defined by the terciles of the inverse elasticity in each country. If we rank goods by market power, we find that the median estimate for a low market power good in Russia is 0.5 with an associated confidence interval of [0.2,0.7]. The corresponding values are 1.8 and [0.8,3] for the median medium market power good and 33 and [3,53] for the median high market power good. Thus, our estimates are sufficiently precise to statistically distinguish the median good in the low, medium and high market power groups.

Obviously, the confidence intervals for a particular good may not be representative of those for all goods in a category. Therefore in Table 3B we report the typical confidence interval, lower bound and upper bound in each category to describe the range of bootstrap estimates.[19] We will be conservative and try to distinguish only between low vs. medium or high market power goods. The data clearly allow us to

[18] This implies calculating more than 3 million bootstrapped parameters. The results were similar when we moved from 50 to 250 bootstraps, which indicates that further increases in the number of repetitions should not change the results.

[19] More specifically, the lower bound of the confidence interval reported is the median lower bound over all the individual confidence intervals estimated in the relevant part of the sample. Similarly for the upper bound.

distinguish between these goods. For example, China's typical *upper* bound for low market power goods is 0.8 whereas its typical *lower* bound for medium or high market power goods is 1.5. This lack of overlap is typical for the sample as a whole where the corresponding values are 0.6 and 1.1. Thus, as we move towards our econometric analysis of tariffs and inverse elasticities, we will be able to use a categorical classification of goods, into low versus medium or high market power, as an instrument to explicitly address measurement error.[20]

4.3 Assessment of Elasticity Estimates

We now turn to the question of whether our estimates themselves are plausible. We do so by first discussing whether their magnitudes are reasonable and then by testing if their variation across countries and goods, which is what we explore in the tariff estimation, is the one predicted by theory and conventional wisdom.

Consider first the magnitudes of the elasticity estimates. Given the absence of alternative estimates, it is difficult to make definitive statements about the reasonableness of the magnitudes we find. One of our interesting findings is that even small countries have market power. This may seem surprising if one assumes the world is composed of homogeneous goods that are traded at no cost. However, this may not be the right framework for thinking about trade for two reasons. First, as we will see, although countries have almost no market power in several homogenous goods, these goods make up only about 10% of the tariff lines in the sample. About 60% of the HS4 goods in the sample are differentiated according to the classification in Rauch (1999) with the remaining 30% classified as reference priced. This may arise because more differentiated goods are, almost by definition, easier to identify and tax differently at the border, hence there are more tariff lines for these goods. Moreover, as we will argue below there is also a good reason to expect that countries have higher market power in differentiated goods and therefore adjust their tariff schedules accordingly.

The second reason why "small" countries can have market power is that trade costs can strongly segment markets. Empirically, we know that these costs have enormous impact on trade patterns and trade volume falls off quite rapidly with distance.[21] These costs also imply that some goods are only traded regionally so that even countries that are small from the world's perspective may have considerable amounts of *regional* market power. For example, Ecuador may represent a large

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[20] The bootstrap estimates may also reflect potential heterogeneity of elasticities across exporters since they are obtained by resampling with replacement. To the extent that they do the resulting measurement error from the heterogeneity is similarly addressed by the instrumenting approach.

[21] According to Anderson and van Wincoop's review of the literature, "the tax equivalent of 'representative' trade costs for industrialized countries is 170 percent" (2004, p. 692). Estimates from gravity equations imply that trade with a partner who shares a border is typically over 14 times larger than with an identically sized non-bordering country if one considers the decay due to distance alone (c.f. Limão and Venables, 1999).

share of demand for certain regionally traded goods, such as Chilean cement, and it is this elasticity that we estimate. This also suggests that countries in regions that are more distant from most of the world's demand have a larger range of products that they only trade among each other. So below we test the hypothesis that importers in more remote regions have more market power.

In practice, the precise value of our elasticities is not critical in determining whether market power affects tariff setting since what we rely on is the variation of market power across countries and goods where we often explore the ranking of goods market power rather than its level. To explore how reasonable the variation in elasticities is in these dimensions we proceed in four ways. First, we check whether elasticities for the same good estimated using data from different countries are correlated. Second, we investigate a particular type of product characteristic, its differentiation, to assess whether the estimates fit our priors. Third, we ask whether countries have more market power when they are larger, as often stressed by the trade literature, or in more remote regions. Finally, we estimate the implied pass-through rates of tariffs to domestic prices and compare them to others in the literature.

The motivation for the first three tests is clearer if we note that the residual supply of exports faced by importer i, m^*_{igv}, is by definition the difference between the production of good g in country v and any consumption in countries $j\neq i$. The export supply elasticity faced by i, $1/\omega_{igv}$ is generally an increasing function of both the exporter's production elasticity, denoted by λ^*_{gv}, a weighted average of demand elasticities, $\tilde{\sigma}^*_{j\neq igv}$; and a decreasing function of i's import share, m^*_{igv}/m^*_{gv}. We discuss the tests in terms of the inverse elasticity, ω_{igv}, so we summarize the relationships above as

$$\omega_{igv} = \omega(\underset{-}{\lambda^*_{gv}}\ ,\ \underset{-}{\tilde{\sigma}^*_{j\neq igv}}\ ,\ \underset{+}{m^*_{igv}/m^*_{gv}}\ ) \qquad (12)$$

Let us first examine whether we obtain similar export elasticities for a given good with different datasets. While it is clear from equation (12) that these elasticities can vary across importers, it is also clear that some goods may be more elastically supplied than others for all importers. The reason is simple: the export supply curve faced by any two importers of a given variety, i.e. from a given exporter, shares at least one common term, the value of the exporter's production. Thus ω_{igv} and ω_{jgv} both reflect the same production elasticity, λ^*_{gv}. Moreover, to the extent that countries other than i and j also consume the good then ω_{igv} and ω_{jgv} also both reflect those consumption elasticities. If our estimates are reasonable then this relationship should also be reflected in the "average" elasticities over exporters of a given good, ω_{ig} and ω_{jg}. Thus, for each country i we regress $\ln(\omega_{ig})$ for all its goods against the mean of $\ln(\omega_{j\neq ig})$ computed using the data of the remaining 14 countries. We report these results in Table 4. The point estimates are all positive and significant, which indicates a very strong positive statistical relationship, with a t-statistic of more than 9 for the typical country. Since the datasets are completely different and each elasticity was estimated independently, these results show that our measure of market power contains information about systematic variation across goods.[22]

We now ask what product characteristics drive the result above and whether they fit our priors. Our second test addresses this question by focusing on product differentiation. As we note in equation (12) we expect countries to have lower market power in goods with higher elasticity of substitution in consumption, e.g. commodities. The reason is simple, if China decreases its demand for a commodity, e.g. U.S. soybeans, and as a consequence their price falls, then other countries will substitute towards that good and away from other sources of supply (e.g. demand less of Brazilian soybeans and other types of beans) so the equilibrium price decline will be minimal. Such substitution is much less likely for specialized or differentiated goods such as locomotives, aircraft or integrated circuits because these are more likely to be tailored for particular markets. Thus we conjecture that countries have more market power in differentiated goods than commodities.

Rauch (1999) classified goods into three categories – commodities, reference priced goods, and differentiated goods – based on whether they were traded on organized exchanges, listed as having a reference price, or could not be priced by either of these means. Table 5 uses this classification and confirms the prediction by testing the differences of the median and mean market power across these categories. The ranking is exactly as expected with the highest market power in differentiated goods followed by reference priced and then commodities. The most striking feature of the table is that both the median and the mean market power are significantly higher for differentiated products– its median value is 2.4, which is about 3 times larger than reference goods and 5 times the value for commodities. This pattern is also clear when we look at the median in each category for individual countries, as shown in Figure 2.

We find a similar pattern if we look at specific goods. For example, among the set of goods with the largest import shares in this sample, the three goods with the least market power are soybeans, barley and natural gas, all with inverse elasticities below 0.1. All of these are commodities for which it is reasonable to expect that a single importer would have a small impact on world prices. In contrast, the median market power in goods such as printed books, locomotives and integrated circuits are more than double the sample median. These are all differentiated goods for which it is more likely that even

[22] We use a log specification to minimize the influence of the outliers. The other motive for using the log specification is that the estimation procedure for the elasticities cannot yield non-positive estimates. Thus the distribution of estimates is skewed with positive deviations from the median vastly exceeding negative ones in magnitude. However, the density function of the log of the inverse export elasticity estimates has a pattern quite similar to a normal density plot.

a single importer can have market power. Thus our methodology generates a reasonable ordering for major import categories.

As a third check for the "reasonableness" of the elasticities, we examine whether they reflect the common intuition that market power increases with country size. Since the sub-sample of products for which we can compute elasticities differs somewhat across countries, computing simple means and medians across different sets of goods may be misleading. Thus, we include HS 4-digit dummies in the regression so as to compare market power for different countries within each import good. The first column in Table 6 reports the results from the regression of log inverse export elasticities on log GDP. There is a positive relationship, which supports the notion that market power rises as GDP rises.[23] This is true even after controlling for the clustering of the standard errors. Although GDP is often strongly positively correlated with import shares the latter are more appropriate for the current purpose, as noted in equation (12). We also obtain a positive relationship when we use an importer's market share in each good instead of GDP. Moreover, this remains true even if we drop China. Hence our estimated elasticities also pass our third "reasonableness" check – larger countries have more market power.[24]

When trade costs are sufficiently large, some goods are only traded regionally. This suggests that for any given GDP, a country in a more remote region would be expected to have higher market power as it accounts for a larger fraction of the region's demand, i.e. it has a larger value for m^*_{igv} / m^*_{gv} in (12). We confirm this in the second column of Table 6 by including a standard measure of remoteness – the inverse of the distance weighted GDPs of other countries in the world.

A final assessment of our elasticities is to consider their implied pass-through rates, i.e. the fraction of the tariff factor increase that is passed-through into higher domestic prices. We compute these rates and compare their magnitude and variation over goods to those in the literature. Unfortunately, few studies estimate the *tariff* pass-through. However, there is an extensive literature on the exchange rate pass-through. We can draw on the latter because the two pass-through rates should be identical in a number of cases, a hypothesis confirmed for example by Feenstra (1989).

We calculate the pass-through, ζ_{ig}, as the effect on the domestic price in country *i* of a given good *g* of a one percent increase in *i*'s tariff factor, i.e. one plus the tariff rate, on all exporters of *g*.

[23] This is consistent with the results in Markusen and Wigle (1989) who use a CGE model to calculate the welfare effects of scaling up all baseline tariffs and find a larger optimal tariff for United States than for Canada.

[24] When we include both the GDP and import share measure we obtain positive coefficients for both but the import share variable is not significant. Although this is partly due to their correlation, the small amount of variation explained by the import share (shown by the R-square within) implies that one must be careful about using it as a proxy for market power. It is possible that regional import shares are a better measure but given the low R-square we doubt that they would be fully satisfactory as a proxy either. The within R-square for GDP is also small, which explains why tariffs and GDP in our sample do not have a robust positive correlation (e.g. it disappears once we drop China) but tariffs and inverse elasticities do, as we show in the next section.

Since $p_{gv} = (1 + \tau_g)\, p^*_{gv}$, we have $\zeta_{ig} = 1 + \mathrm{d} \ln p^*_{gv} / \mathrm{d} \ln(1 + \tau_g)$. In the theory appendix, we show that in our framework the last term is simply $-\omega_{ig}/(1 + \omega_{ig})$ and so $\zeta_{ig} = 1/(1 + \omega_{ig})$. Thus using the median inverse elasticity of 1.6, the typical pass-through rate in our sample is 0.4, which is similar to the values in the literature. Kreinin (1961) finds a pass-through of about one third for U.S. tariff reductions. Chang and Winters (2002, p. 898) find imperfect tariff pass-through for Brazil's imports from Korea (0.18), Germany (0.74) and U.S. (0.89). The survey by Goldberg and Knetter (1997) reports that the typical *exchange* rate pass-through is 0.6 with large variation across industries (p. 1250).[25]

Finally, consider the variation of the pass-through across types of goods. Using the median estimates in Table 5 we find that it is highest for commodities (0.7) then reference priced goods (0.6) and lowest for differentiated products (0.3). For specific commodities such as barley, soybeans and crude oil we find nearly full pass-through. Thus these estimates are consistent with our priors and with studies of exchange pass-through that find higher rates for commodities than other goods.[26]

In sum, the analysis above suggests that our elasticity estimates are "reasonable" by a number of criteria. We now ask if they are an important determinant in setting tariffs.

5. Estimating the Impact of Market Power on Tariffs

5.1 Preview

We can now answer whether there is any relationship between the tariffs and export supply elasticities. Before turning to the regression evidence, we will examine a data plot: the median tariff in each country against the median inverse export elasticity. There are many reasons to be skeptical that we can obtain a relationship in the country cross section. We only have fifteen countries to work with so one may worry that any one country can dominate the results. In addition, the countries in our sample have very different political systems, economic conditions, and mix of other protectionist tools – all of which are reasons to abandon all hope that a relationship will be visible. However, since the cross-sectional story of market power and tariffs has such prominence, it is worthwhile examining it.

Figure 3 shows that there is a strong positive relationship between the median tariff in a country and market power in the typical good, as measured by its median inverse elasticity. The pattern does not seem to be driven by any one country or even set of countries on a particular continent or with a

[25] Note that our pass-through measure does not reflect the import substitution across *varieties* of g, σ_{ig}, because we are considering an increase in the tariffs on all exporters of that good. The elasticity of substitution across *goods* in the model in the appendix is unity. This also highlights a difference between our approach, where imperfect pass-through can occur even if the exporter has no market power, and the one in most pass-through studies. The latter typically posits an imperfectly competitive exporter (facing an elasticity of demand greater than unity) and then estimates a reduced form price equation obtained from its optimal decision to adjust markups (or costs) in response to changes in the exchange rate or tariff factor.

[26] Campa and Goldberg (2005) for example estimate that the average in OECD countries for raw materials is 0.6 whereas it is 0.4 for manufacturing (p. 690).

particular income level. The positive relationship between median tariffs and median elasticities is also statistically significant.[27] Of course, there are many reasons to be wary of this relationship, as we just pointed out. Fortunately, the vast quantity of country-good data underlying this plot can be used to examine the relationship more carefully and in our working paper we confirm its robustness.[28]

The result we have presented thus far is suggestive but still far from convincing. Expressing the optimal tariff purely in terms of a country's characteristic such as size and resulting market power may be natural in a two-good model, but is not very useful from an empirical perspective because of the many cross-country differences that may affect average tariff levels. Furthermore, as we have seen, the theory also provides important predictions for tariff variation within a country. Since there is considerable variation in tariffs and elasticities within countries and fewer potential omitted variables our main results in the next section follow this route.

5.2 Baseline Results

Our approach to estimating the impact of market power on tariffs is two-pronged. In this section we provide baseline results from specifications where the inverse export supply elasticity is the key determinant of protection, and we include country and industry effects to control for tariff motives highlighted by various political economy models. In the next section, we further augment the model to include two specific prominent motives for protection: tariff revenue and lobbying. The baseline estimates are robust to either of these, both qualitatively and quantitatively. Given this and the fact that we do not have the required data for all countries for the augmented model we first present the baseline results.

The general econometric model we employ can be written as follows

$$\tau_{ig} = \beta_i f(\omega_{ig}) + \boldsymbol{\eta}_{iG} + \mathbf{x}_{ig}\boldsymbol{\gamma} + u_{ig} \tag{13}$$

where the ad valorem tariff, τ, varies by country *i* and HS 4-digit good, *g*, as does the market power variable, *ω*, and *G* defines the industry of good *g*. Although the basic theory yields a linear relationship, we have discussed theoretical reasons to expect the true effect to diminish at higher levels of market power. Because of this and of econometric reasons, we also consider alternative functional forms for *f*(.). Since our main objective is to establish whether market power is a significant determinant of

[27] If we regress the median tariff on the median inverse elasticity we obtain a positive slope (b=5.9; s.e. = 2.9; R^2 = 0. 21). The positive relationship is still present if we exclude China, (b=4.2; s.e. = 2.36).

[28] In the original working paper version (Broda et al, 2006) we report the results of regressing average tariffs on the inverse export elasticity controlling for HS-4 digit fixed effects. This uses the variation within a product and across countries and confirms that the result in Figure 3 is robust to issues of measurement error, endogeneity and censoring.

tariffs rather than to establish in which countries the marginal effect is stronger, the baseline results focus on the case where $\beta_i = \beta$ for all countries. However, we also present country specific regressions.

As we discuss in the theory section, the tariff may depend on various other factors. Some are country specific, e.g. country location, level of development, expected WTO accession, formerly communist, etc. Several others depend on political economy factors that are not easily observable. However, many of the latter factors are channeled by lobbies organized at the sector or industry level, where each industry, G, includes a different subset of goods, g. A flexible way to capture the impact of such determinants on tariffs is to include country and/or industry effects. Therefore we consider three alternatives. First, estimating only the country effects and treating any industry-country factors, v_{iG}, as part of the error term, so that in (13) we have $\boldsymbol{\eta}_{iG} = \boldsymbol{\eta}_i + v_{iG}$. Second, including country and common industry effects, i.e. $\boldsymbol{\eta}_{iG} = \boldsymbol{\eta}_G + \boldsymbol{\eta}_i + v_{iG}$ for all i. This controls for the fact that there is considerable variation in trade protection across industries. However, any given industry can have very different levels of protection across countries and therefore, the most general case is one where $\boldsymbol{\eta}_{iG}$ represents a set of industry-by-country effects. The latter is the case we mostly focus on since it controls for a considerable amount of unobserved industry heterogeneity and allows us to identify the effect of market power on tariffs by exploring product variation within countries and industries.

Most theoretical trade policy models focus on industry level determinants, which we will be able to control for as just described. However, in any given country there may exist certain product characteristics that are correlated with market power and affect the tariff set in that product. There is little empirical guidance on what these other relevant characteristics are (since most studies are conducted at the industry level) and it is therefore impossible to ensure that all relevant ones are controlled for. Thus our main strategy for addressing omitted product variables in this section is to use instrumental variables. In the next section we test if the IV results are robust to controlling for some key determinants of tariffs represented by the vector $\mathbf{x}_{ig}$ in (13).

Table 7 presents OLS and Tobit estimates for the pooled sample. The first three columns include country effects. The next six columns also include industry effects, which are defined according to the 21 sections of the Harmonized Tariff Schedule, e.g. textiles, chemicals, plastics, etc. Since the results in columns 1-3 are qualitatively similar to the comparable ones in columns 4-6, we discuss the latter.[29] The linear market power measure, in column 4, has a positive and significant effect on tariffs. The coefficient is small because of a few large outliers in the inverse elasticity, as we previously discussed. Moreover, the effect represents an average of increases in the market power at

[29] The comparable three specifications with industry by country effects are also similar and available on request.

low and high levels. When market power is high, the tariff is closer to being prohibitive and the marginal effect of further increases in market power are expected to be small. This is confirmed in column 8 by a regression where the knot for the different slopes is endogenously determined by the data. Despite the lower marginal effect at high market power those goods do have significantly higher average tariffs.[30]

A parsimonious way to address the skewness of market power and its non-linear impact on tariffs is a semi-log specification, i.e. to use $f(\omega) = \ln(\omega)$ in (13). The OLS estimate in column 6 shows that market power also has a positive and significant effect on tariffs under this specification. The result is identical for the Tobit specification in column 7 where the tariff censoring point is zero.

In column 5 we address the measurement error and, to some extent, the functional form issue, by sorting each country's data by the inverse export elasticity and creating a dummy equal to 1 if it is above the 33rd percentile. The estimate shows that goods with higher inverse elasticities have higher tariffs. In column 9 we find that this difference in tariffs relative to goods with low market power is similar for goods where market power is high (above 66th percentile of inverse export elasticity) or medium (between 33rd and 66th). This confirms the diminishing marginal effect we found and further supports the use of a flexible functional form such as the semi-log or dummy.

The OLS estimates are potentially biased because of attenuation caused by measurement error and omitted variables. Since our objective is to determine causality and provide a quantification of the effect of market power on tariffs we must address these sources of bias. In this section we do so by using instrumental variables.

The main instrument we employ for a given country's market power in a good is the average market power in that good in the other countries. The basic motivation is simple: to minimize the country-product specific portion of market power that may be correlated with other determinants for the tariff on that good in a particular country. In the semi-log specification we could use as an instrument the average of other countries' log inverse elasticities since the variables are strongly correlated, as shown in Table 4. This procedure addresses endogeneity concerns. However, when we employ the continuous measures, this procedure alone does not necessarily eliminate the measurement error since if there are at least a couple of countries with large measurement error in any given good, the instrument itself will have considerable error. When assessing the elasticities we showed that our estimates clearly distinguish between goods where a country has low vs. medium or high market power.

[30] It estimates a slope of 1.9 when market power is below the estimated threshold (53rd percentile), which is considerably larger than the slope above it. The threshold in a similar specification without industry effects is at the 33rd percentile.

Thus in equation (13) we instrument $f(\omega_{ig})$ with the average of the categorical variable for all countries other than i. This instrument is also strongly positively correlated with $f(\omega_{ig})$.

Table 8 presents the IV results for the pooled sample of 15 countries. The results for the full sample in Panel A contain three specifications: level, dummy and semi-log. We estimate each including country effects (columns 1-3), country and industry effects (4-6) and industry-by-country effects (7-9). The latter specification best isolates the impact of market power, but comparing it to the other specifications also provides useful insights.

The first point that stands out is that the estimated market power effect on tariffs is positive for all specifications and considerably larger than with OLS. In the semi-log specification for example the coefficient is 1.7 when we control for industry or industry-by-country effects in columns (6) and (9) respectively. This estimate is ten times larger than the OLS one and significant at the 1% level.[31] The dummy estimates in columns 5 and 8 illustrate a similar point. Products in which countries have medium or high market power have tariffs about 9 percentage points higher, a result that is both economically and statistically significant. Since the dummy is less prone to measurement error, these results suggest there was a downward bias due to omitted variables that is addressed by the IV. We will thoroughly discuss the magnitude of these effects in section 5.6.

A third point worth noting is the importance of accounting for unobserved industry heterogeneity when we employ a parsimonious specification. The estimated market power coefficients in columns 1-3 generally double after we account for such heterogeneity in columns 4-6 and 7-9.

The linear version is unlikely to be the correct functional form, as both the data and basic extensions of the theory strongly suggest. However, given its prominence in the basic theoretical prediction we also present baseline results for it. The more general specification in column 7 confirms the results obtained with the semi-log and dummy: a positive and significant effect that is considerably larger than the OLS estimate.

Recall from the data preview section that Bolivia, Oman and Saudi Arabia had little variation in their tariffs with most grouped in two or three value bins. A linear regression approach is generally not the most appropriate way to treat these observations – an issue we address below when we estimate the regressions for each country. Therefore one may argue that a more accurate estimate for the pooled sample should exclude these countries. These results are presented in panel B of Table 8. The key difference relative to the full sample is an increase in precision (in terms of the z-statistics). There is

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[31] There is also indirect evidence that our IV approach addresses the measurement error in ω satisfactorily. Recall that this was most important for estimates above the 90th percentile in each country. However, when we re-estimate the IV without those observations we obtain very similar estimates for β.

also an increase in the magnitude for the dummy and semi-log specifications. The high partial *F*-statistics from the first stage regression, in the last row, show that the instrument performs well.

5.3 Individual Country Results

To carefully establish the tariff determinants of any given country requires its own paper. However, we want to determine whether the baseline results represent trade policy setting in the typical country. We remain as close as possible to the framework we have used so far. However, we cannot ignore obvious issues such as the bunching of tariffs in Bolivia, Oman and Saudi Arabia. For the other 12 countries we still employ the IV approach with industry effects and estimate the unrestricted version of (13) for each country.

Tariffs in Bolivia, Oman and Saudi Arabia have little variation and almost none within industries. So it is doubtful that we can find a strong relationship for them and it is clear that we require a different econometric approach to address the fact that a large fraction of their tariffs appear to be censored below and/or above. Thus for these three countries we run censored regressions where we also instrument for the market power variables.[32]

Table 9 presents the IV results by country. Panel A focuses on the semi-log specification. The first two columns reproduce the pooled results from Table 8 for ease of comparison. The estimate is positive for each and every one of the 15 countries. It is also statistically significant at the 5% or 1% level for all but two..[33] The estimate for the typical country is 1.75 and the mean is 2.15. These are very close to the pooled estimates, which were 1.73 for the full sample and 2.11 for the sub sample of 12 countries.

A measure that is more directly comparable across countries is the implied elasticity of tariffs with respect to market power. We obtain it by dividing the coefficient by the mean tariff and show it in the last row. This value ranges from 0.13 to 0.15 in the pooled estimates. The value is similar to the mean over the country estimates, 0.17, as well as the value obtained for the typical country, 0.16, both of which are shown in the last two columns. The range of elasticity estimates across countries is fairly narrow, from 0.08 to 0.23 for 13 of the countries, and it does not have an obvious pattern. So the pooled estimates capture an effect that is typical of the countries in the sample.

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[32] Most tariffs in Bolivia for example are set at a maximum of 10, so we run an instrumental variable tobit with that value as an upper limit. In Oman 10% of the tariffs are set at 0 and nearly all others at 5% so we use two censoring points. For Saudi Arabia about 87% of observations have tariffs equal to 12 with most others above it so we use that as the lower limit.

[33] One of them is Saudi Arabia, where we did not expect a precise estimate anyway. The other is the Czech Republic that, as we note in Table 1, set its tariffs in 1992 as a federation with Slovakia. Since this federation was a member of GATT its tariffs are less likely to reflect a terms-of-trade motive, possibly explaining our finding here. Note also that the pooled results in tables 7, 8 and 10 are robust to dropping the Czech observations and available upon request.

The pattern of heterogeneity in the point estimates across countries in Table 9 provides some additional support for the theory. Countries are sorted in decreasing order of their 1996 GDP. The largest, China, has the highest coefficient. More generally, the larger economies tend to have larger estimates. When we test this directly we find that the difference in the estimate is large and significant.[34] However, as we pointed out in the previous paragraph there is no such pattern in the *elasticity* of tariffs with respect to market power, that is after we divide the estimated coefficient by the average tariff. The reason for this is that in this sample average tariffs are higher for larger countries, as the theory would predict.

Panel B presents the analog using the categorical variable. They are qualitatively similar to panel A – all positive and also all significant, except for two countries. The typical country sets tariffs that are 9 percentage points higher in goods with medium or high market power relative to those where it has low power. This effect is almost the same magnitude as the average tariff for the typical country, as we can see in the last row. This is a preview of the large effects implied by the terms-of-trade motive that we analyze in detail below.

5.4 Other Robustness Tests

By construction, much of the variation in our instrument is across goods. This is one key reason to focus the analysis within countries and across goods. In the pooled regression the fact that the instrument varies mostly across goods could induce a correlation in the error term for any given good across countries. Clustering the standard errors by HS-4 addresses this concern and we verified that it does not change the significance of the results.

We can relax the assumption of common elasticities across sets of exporters of a given good to country *i*. However, this is costly since it reduces both the number of elasticities we can estimate and their precision. It also raises the question of how we should aggregate the elasticities over exporters of the same good to estimate the tariff equation, since the tariff data we use for *i* does not vary by exporter. Nonetheless we test if the results are sensitive to the sample of exporters used. Effectively we calculate two estimates for *each* ω_{ig} each using a different set of exporters. For each importer *i* we rank exporters by their total export value in *all* goods shipped to *i* in the entire period. The even sample includes exporters with even ranking and the odd includes the remaining. We then re-run the baseline results in columns 7-9 in Table 8 with each set of estimates. Both sets yield a positive and significant

[34] That is, in the general specification we model $\beta_i = \beta + \beta^L * 1(\text{Size}_i)$, where the indicator variable 1(Size) is one if country i is above the 66th percentile in terms of size. We defined size as either GDP in 1996 or GDP adjusted by "regional market size", i.e. divided by a distance weighted average of other countries' GDPs, both yield the same 6 countries. The additional instrument required is simply the original instrument interacted with $1(\text{Size}_i)$. The estimates for β remain positive and significant and the extra effect for the larger countries was 2.6 for the semilog (s.e. 0.5) and 12 for the dummy (s.e. 2.3).

effect of market power on tariffs. Moreover the quantitative result is nearly identical across the sets. So the precise selection of exporters in estimating the elasticity does not change the key finding.[35]

5.5 Augmented Models: Revenue and Lobbying

The preceding analysis established that these countries set higher tariffs in goods in which they have more market power. However, trade policy can also be strongly influenced by revenue considerations and domestic political interests. It is hard to see why there might be a systematic correlation between our estimates of export supply elasticities and political economy variables because all importer-industry-time variation has been purged from the elasticity data before estimation. Moreover, lobbies tend to form at the industry level and, as we reported above, our results are stronger when we include industry effects that account for unobserved heterogeneity. Finally, the IV estimates indicate that the results are strengthened when we address potential endogeneity problems such as the one that could result from omitted variable bias.

Nonetheless, we want to further test if our baseline results are biased due to omitted variables. This would only occur in our IV estimates if certain product characteristics that affect tariffs are also correlated with our instrument for market power. As noted before much of the theory and empirical evidence on trade protection focuses on industry-level determinants. So there isn't an obvious list of *product* characteristics we should control for. Thus we focus on two prominent motives for protection that also have implications for cross-product variation in protection: revenue and lobbying. This also allows us to determine the importance of market power relative to these prominent alternative explanations for tariff setting

Consider first the use of tariffs to collect revenue. To the extent that this motive is correlated with the level of development, it is captured by country effects. However, it is simple to show that if governments use tariffs to raise revenues they would impose higher tariffs on goods with lower import demand elasticity, σ_{ig}. This occurs because when a given tariff rate is imposed on a good with lower import elasticity, it raises more revenue and imposes a lower distortion for the standard Ramsey taxation motive. If the import elasticity were correlated with the foreign export supply elasticity, our results could simply be picking up the tariff revenue motive.

The results in Table 10 address this question. We obtain the inverse import demand elasticity, $1/\sigma_{ig}$, using the procedure described earlier and instrument it using the same approach as the market power variable. That is, we create a categorical variable for each country that is 0 for product *g* if its

[35] We omit the table for space considerations but it is available upon request.

value of $1/\sigma_{ig}$ is in the bottom tercile in country *i* and 1 otherwise. We then use its average for that product over the other countries as the instrument.

In columns 3 and 4 of Table 10, we include the tariff revenue variable to augment the baseline model with industry-by-country effects. The coefficient on the dummy variable for market power is 9 -- identical to the baseline reproduced in the first column. The estimate for the semi-log is also statistically identical to the baseline, 1.8 instead of 1.7. So, once we have accounted for industry-by-country effects and instrumented the market power variable, our baseline estimates do not reflect an omitted variable bias arising from a tariff revenue motive for tariffs.[36]

The specific political economy factors that are relevant for the tariff structure can also differ across these countries. So, we now include a political economy variable that is central in an important model, Grossman and Helpman (1995), and that also plays a role in alternative political economy models. When all sectors are politically organized, the Grossman-Helpman model provides a parsimonious characterization of the effects of both market power and domestic lobbying. In this model, tariffs are given by the sum of the inverse elasticity and what we refer to as the lobbying variable, z_{ig}/σ_{ig}, as defined in equation (7) with I_{ig}=1.

Recall that the variable z_{ig} is the ratio of domestic production value to import value, where the latter excludes tariffs. Thus it requires production data, which we could obtain for 7 of the 15 countries in our sample for years close to the tariff data.[37] This is only available for all these countries at the ISIC 3-digit data from UNIDOs industrial database. So z_{ig} can be interpreted as country *i*'s average penetration for the goods in that ISIC 3-digit category. Since we divide this by the import demand elasticity, which varies by HS-4, the lobbying variable also varies at the HS 4-digit level.

In the regressions, we treat the lobbying variable similarly to market power. More specifically, we employ either its log or a categorical variable that takes the value of 0 for the lower tercile of z_g / σ_g in that country and 1 otherwise. We instrument the variable since production and imports depend on tariff levels. The instrument is constructed by taking the average of the categorical lobbying variable over the remaining countries for each good. As indicated by the partial *F*-statistics for the first stage in Table 10, the instrument used is strongly correlated with the lobbying variable.

The last two columns of Table 10 present the estimates when we augment our baseline estimates with industry-by-country effects using the lobbying variable above. The market power effect in the dummy specification is 10 percentage points – larger but statistically indistinguishable from

[36] This result is robust to dropping Bolivia, Oman and Saudi Arabia, the only difference being that the tariff revenue coefficient becomes positive in the dummy specification but still insignificant.

[37] These countries are Bolivia, China, Ecuador, Latvia, Lithuania, Taiwan and Ukraine.

either the baseline for the full sample, 9, or the one in the sub sample of 12 countries, 11. The same conclusion holds if we consider the semi-log specification where the estimate is now 1.9.

Note also that the reason why the results are similar is *not* because we are adding an irrelevant variable. Several studies found that a similar variable is empirically important for other countries and we find that it is significant for this sample as well. Below we quantify the importance of market power in tariff setting not just by itself but relative to this important alternative explanation.

Table 11 shows that these results are not driven by a single country. In all 7 countries for which we have production data, market power has a positive effect on tariffs which is statistically and economically significant. We now turn to a more detailed quantification of the effects.

5.5 Quantification

Table 12 provides summary statistics for the key parameters estimated and computes simple counterfactuals that illustrate the economic and statistical importance of market power in tariff setting. The columns correspond to the main specifications in the baseline and lobbying augmented models. The results are fairly similar across specifications, so we focus on the pooled results with the lobbying variable in which we can assess the relative importance of the terms-of-trade motive.

The first row reproduces the coefficient for the semi-log specification. In the case of the pooled regression in Table 10 it indicates that a one log point increase in market power increases tariffs by 1.9 percentage points. Thus one standard deviation in market power leads to a 5 percentage point increase in tariffs, which is large if we recall that the median tariff in this sample is 10 percentage points. The market power effect is also important relative to the lobbying variable. Since their coefficients are similar so is the elasticity of tariffs with respect to these two important determinants.

Alternatively, we can ask how important market power is in explaining tariff variation. We will do so with respect to a natural benchmark: the importance of political economy. To do this, we compute the impact of a standard deviation in each of those variables on tariffs. This effect is about 1.6 times larger for market power, indicating that it is more important in explaining tariff variation in these countries than the lobbying variable. We obtain a similar value if we use the dummy variable, as shown in the last row.

To gauge the economic importance of the terms-of-trade motive relative to all tariff setting motives we can compare the average effect due to market power to the average tariff. The effect ranges from 8 to 11 percentage points, depending on the specification. For the typical country this terms-of-trade motive is about the same magnitude as their average tariff, more specifically it is 92% of the average in the baseline and 97% in the lobbying augmented model. Given that the effect applies to two-thirds of the tariff lines the implied tariff reductions if these countries did not exert their market power would be substantial, much larger than the 25 percent target reduction for developing WTO members in the Uruguay Round.

One important motive for our study is that if countries possess and exert market power then their tariffs depress the price received by foreign exporters. Thus, reducing tariffs increases those prices. Our estimates allow us to calculate the price increase for exporters when an importer, *i*, reduces or eliminates its use of tariffs for a terms-of-trade motive. This is an interesting counterfactual because this type of beggar-thy-neighbor trade policy is believed to have been central in the trade war of the 1930's and arguably one of the key motives behind the creation of the GATT. In fact, Bagwell and Staiger (1999) suggest that when a country enters a trade agreement such as the GATT/WTO its tariffs no longer reflect the market power component (in a 2-country symmetric model) or reflect it only partially. The counterfactual also provides an estimate of the impact of a country leaving the agreement and re-exerting its market power, which is obtained by simply by reversing the direction of the exporter price effects.

The price received by exporting variety v of good g to country i can be written as $p^*_{igv} = p^*_{igv}\left(\tau_{ig}(\omega_{ig}),.\right)$, i.e. a function of the tariff it faces and other parameters that are omitted. The percent increase in that price as an importer starts to treat a given good with medium or high market power as if it had low market power is then

$$\text{(14)} \qquad \begin{aligned} \Delta \ln p^*_{igv} &= \left.\frac{\Delta \ln p^*_{igv}}{\Delta \ln\left(1+\tau_{ig}\right)}\right|_{\omega_{ig}^{med\_hi}} \times [\ln\left(1+\tau_{ig}\right)\Big|_{\omega_{ig}^{low}} - \ln\left(1+\tau_{ig}\right)\Big|_{\omega_{ig}^{med\_hi}}]\times 100 \\ &\approx -\left(\zeta_{ig}-1\right) * \beta_i \end{aligned}$$

where ζ_{ig} is simply the domestic pass-through, which we discussed in our assessment of the elasticities, and so ζ_{ig} -1 is the effect of the tariff factor on exporter prices. In equation (13), β_i measures how much higher tariffs are in a good with medium or high market power relative to low.[38]

In the full sample the typical good in the middle tercile in terms of market power has an inverse elasticity of 1.6. Therefore, ζ_{ig} -1 = -0.6 (= 1/(1+ 1.6) -1). Multiplying this by the estimated tariff reduction due to treating those goods as low market power goods (which we estimate to be 9 percentage points in Table 8, column 8) we obtain a 5 percent increase in the price received by the

[38] The approximation arises because in equation (13) we use tariffs rather than ln (1 + τ_{ig}), but these are identical for most tariffs in the sample. When they differ, e.g. China, we report the results with the tariff factor. Note also that the estimate for ζ_{ig} -1 assumes that all exporters of good g face similar tariffs. In the WTO interpretation of the counterfactual this implies they would all be in the WTO. Otherwise the effect would be larger since the tariff reduction on exporter v alone would lead to a additional import demand for v due to substitution away from exporters of other varieties of g not in the WTO.

exporters. By construction, this applies to 1/3 of goods in each country. For the typical good in the high market power bin we have $\omega_{ig} = 54$ (Table 3A), and thus any given tariff reduction has a larger effect on export prices. In this case the change in exporter prices for these goods is close to 9 percent.

Finally, for the three largest countries, the effect is stronger than the average. Using the country specific estimates, the effect for Russia and Taiwan is 6-10% and for China it is particularly large, 17-25%. Thus, if entry into the WTO leads these countries to remove the portion of their tariffs driven by a terms-of-trade motive, exporters into these markets will enjoy a large benefit from this direct price effect. By the same token, if these countries were to abandon an agreement and re-exert their market power, this would be quite costly for the foreign exporters facing the higher tariffs.

In sum, the terms-of-trade motive is economically and statistically important for tariff setting in these countries. It is more important in explaining tariff variation than a key political economy variable used in previous studies. Moreover, it causes significant changes to prices received by foreign exporters particularly as they try to sell in the larger countries such as China, Russia and Taiwan.

6. Market power and trade barriers in a large developed WTO member

Our focus on non-WTO members is motivated by the theory, which predicts that a country's tariffs will reflect its market power when it acts non-cooperatively. Since the terms-of-trade gain for the importer is lower than the corresponding cost to the exporter, cooperation between the two, e.g. as they become WTO members, could attenuate or eliminate this motive for tariffs. However, a reasonable question is whether the forces we identify in this paper are present for any instruments of protection used by WTO members. Obviously, we cannot simply analyze if market power affects the MFN tariffs of WTO members because a failure to find such a relationship could simply be due to the effectiveness of this agreement in eliminating the terms-of-trade externality (c.f. Bagwell and Staiger, 1999, and Grossman and Helpman, 1995). Thus we need to consider alternative experiments.

We consider instruments of protection whose levels are not negotiated within the GATT/WTO or are only partially so. While this experiment may not be as clean as our earlier one, it does allow us to explore this question further. Moreover, it can provide insights about how the negotiated trade policies of current members would change if they were not subject to WTO constraints.

We focus on two such instruments for the United States both because it has good data and because it is the world's largest economy. First, we follow most empirical studies of U.S. protection and use non-tariff barriers (NTB's) as the measure of its non-cooperative trade policy (e.g. Goldberg and Maggi, 1999, Bandyopadhyay and Gawande, 2000). Several of these NTB's – e.g. antidumping duties, countervailing duties, and some forms of quotas – generate higher welfare for the importing country if they are implemented in goods where it has market power. Thus the prediction is that NTBs are more prevalent in goods with higher market power. Until recently there were no tariff equivalents of NTBs for a large set of goods. Thus we use the standard measure of NTB strength in the literature: the coverage ratio, i.e. the share of 6-digit goods within each 4-digit classification that contain an NTB. We complement this by using the ad valorem equivalent recently estimated by Kee et al (2006).

Second, we examine "statutory rates" – the tariffs the U.S. applies to non-GATT countries to which it does not grant MFN status. Statutory tariffs are set non-cooperatively, which is apparent from their high levels and the targeted countries.[39] Successive rounds of trade negotiations opened a large gap between these rates and the MFN rates the U.S. negotiates with and sets on WTO members. The average U.S. statutory rate in 1999 was about 30 percent – almost ten times larger than the MFN rate in our sample. Although statutory rates currently apply to a small number of countries, understanding their determinants provides an interesting and unique insight into how the U.S. sets tariffs non-cooperatively. Thus they are a useful complement to U.S. NTBs, which apply to many countries.

Our estimation strategy is similar to the one used thus far. We estimate the relevant elasticities for the U.S. using the same procedure and use them to estimate equation (13) including industry dummies and instrumenting for the remaining variables. The structure of U.S. production, trade, and demand differs in important ways from those of the developing countries we analyzed. Thus we thought it would be more appropriate to instrument for U.S. elasticities and import penetration ratios using data from large developed countries: Canada, France, Germany, Japan, and the U.K. Using this data we construct the instruments as before.[40]

In Table 13 we report the results for the U.S. The results in Panel A show that the U.S. sets significantly higher NTBs in products where it has more market power. This is true if we measure NTBs by the commonly used coverage ratio (columns 1,2,5,6) or the ad valorem equivalent (columns 3,4,7,8). The magnitudes for the ad valorem specification are large since products with NTBs have large tariff equivalents: about 18% for the typical HS-4 with an NTB. These NTBs affect a large number of products – a quarter of the HS-4 lines in the sample.[41] The results are robust to including the

[39] In 1989, the countries subject to these tariffs were Afghanistan, Albania, Bulgaria, Cuba, Czechoslovakia, Estonia, German Democratic Republic, Kampuchea, Laos, Latvia, Lithuania, Mongolia, North Korea, Romania, USSR, and Vietnam. Before 1980 China was also subject to this set of tariffs.

[40] The U.S. elasticity estimates are reasonable when we use the criteria previously applied. First, they are strongly correlated with those of these five countries. Second, the typical inverse elasticity is highest for differentiated (1.6) than reference priced goods (0.55) or commodities (0.41). Third, the pass-through for the typical good is 0.5, which is a common value for the U.S. according to Goldberg and Knetter (1997).

[41] We control for the censoring at zero by using an IV-Tobit. For the coverage ratio we also control for the censoring at one.

lobbying variable, which has a smaller impact on tariffs than market power, as we previously found for other countries.

Panel B focuses on tariff barriers. Market power has a strong and significant positive effect on statutory tariffs, which the U.S. sets non-cooperatively. These rates are between 21 and 27 percentage points higher in goods with medium or high market power (columns 1 and 5). Interestingly, the elasticity at the mean is 0.17, very similar to the typical value we found for the non-WTO countries.

Finally, when we use the U.S.MFN rates (columns 3,4,7,8) we find a much weaker relationship with market power. In fact, although the effect is positive, it is not significant at the conventional levels. Even if we take into account the lower mean of MFN rates we still find a lower elasticity for them than for statutory rates. In sum, the evidence on NTBs, statutory tariffs, and MFN tariffs indicates that market power matters for U.S. trade policy in areas not covered by the WTO. In other words, when the U.S. can set trade barriers non-cooperatively, it takes market power into account. This strongly suggests that market power would play an important role for *all* U.S. trade policy if it were set non-cooperatively, e.g. in the absence of the WTO.

7. Conclusion

The idea that a country can improve its terms-of-trade and welfare through the imposition of tariffs has been in the economics literature for over a century. Since then, economists have known that the optimal tariff is positive for goods that are supplied inelastically. However, no one has tested whether countries set higher tariffs in goods in which they have more market power. This paper is the first to provide evidence that importers who are not members of the GATT/WTO do impose higher tariffs on goods with lower export supply elasticities. This result is present when looking at tariffs across countries, across goods within countries and industries, and even after controlling for tariff revenue and various political economy motivations.

The results go beyond isolating the statistical impact of market power on tariffs. They show that the impact is economically significant as well. It is of the same magnitude as the average tariffs in these non-WTO members and at least as important as the lobbying motive that has attracted much attention in previous work. Thus, removing the terms-of-trade motive for tariff setting would lead to increases in the prices received by foreign exporters to these markets. These increases are significant for the several goods in which we estimate that importers have considerable market power.

We also find that market power strongly affects the non-cooperative trade policies of a large developed country, the U.S. Its statutory tariffs for example are 27 percentage points higher in goods in which it has significant market power. Thus the importance of the terms-of-trade motive extends beyond non-WTO members and understanding its impact on trade policy is essential. The strong effect of market power on non-cooperative policies and its absence on the cooperative tariffs the U.S. negotiates in the WTO indicates a quantitatively important role for this institution in reducing tariffs.

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Figure 1: Tariff Distribution by Country

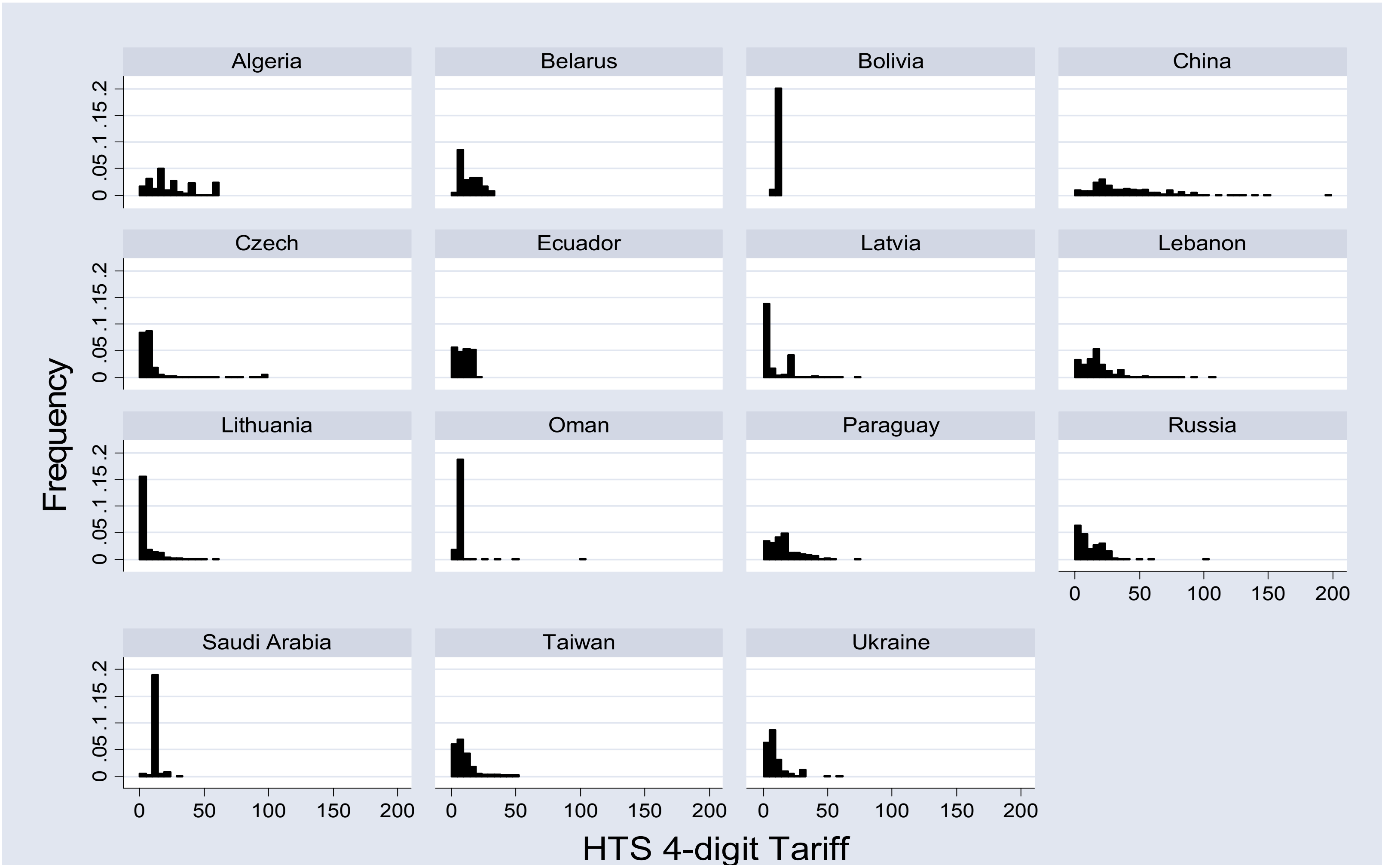

Figure 2: Median Inverse Elasticities by Product Type
(Commodity;Reference Priced and Differentiated Products)

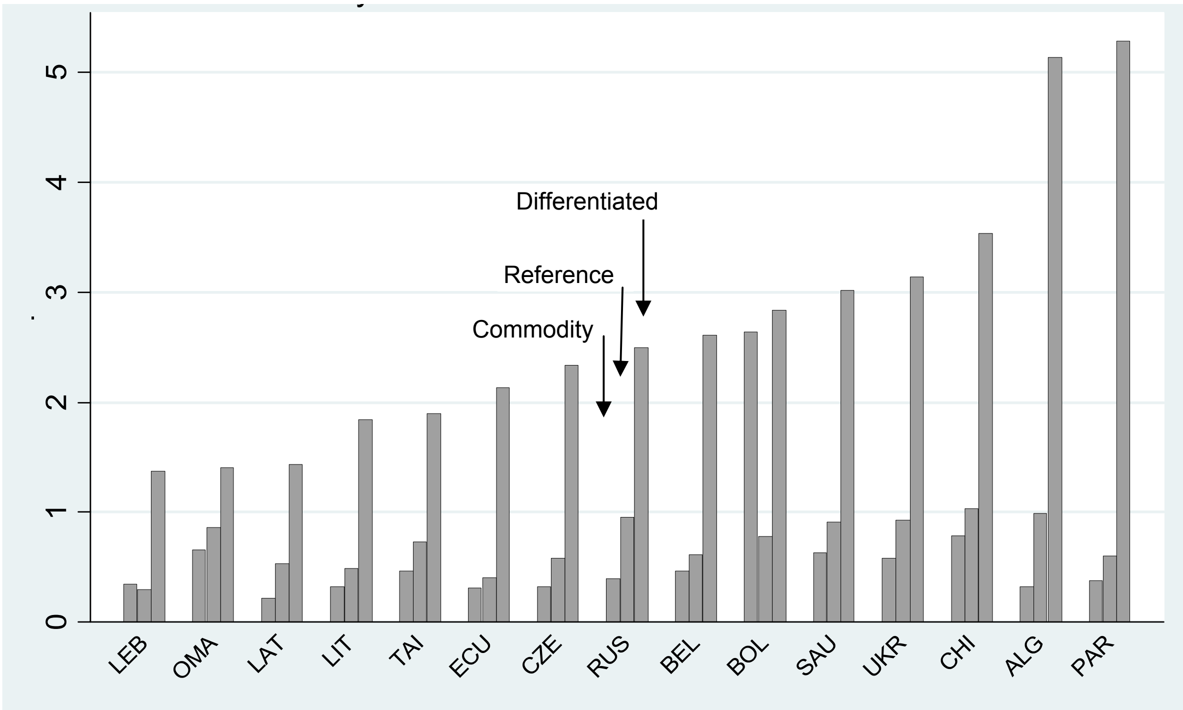

Figure 3: Median Tariffs and Market Power Across Countries

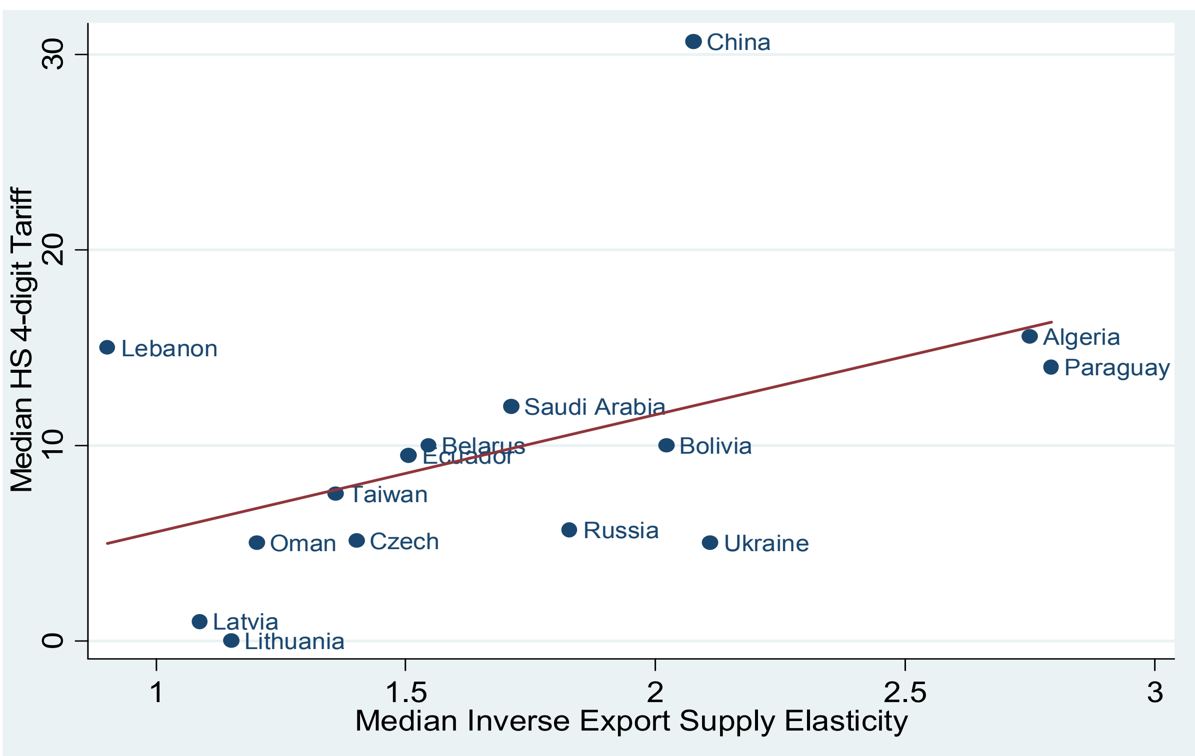

Table 1
Data Sources and Years

	GATT/WTO Accession date	Production Data Source	Production Data Years	Tariff Data [a]	Trade Data [b]
Algeria				93	93-03
Belarus				97	98-03
Bolivia [c]	8-Sep-1990	UNIDO	93	93	93-03
China	11-Dec-2001	UNIDO	93	93	93-03
Czech [d]	15-Apr-1993			92	93-03
Ecuador	21-Jan-1996	UNIDO	93	93	94-03
Latvia	10-Feb-1999	UNIDO	96	97	94-03
Lebanon				00	97-02
Lithuania	31-May-2001	UNIDO	97	97	94-03
Oman	9-Nov-2000			92	94-03
Paraguay	6-Jan-1994			91	94-03
Russia				94	96-03
Saudi Arabia	11-Dec-2005			91	93-03
Taiwan	1-Jan-2002	UNIDO	96	96	92-96
Ukraine		UNIDO	97	97	96-02

Notes: (a) All tariff data is from TRAINS. Countries are included if we have tariff data for at least one year before accession (GATT/WTO). (b) Except for Taiwan, all trade data is from COMTRADE. For Taiwan data if from TRAINS. (c) The date of the tariffs for Bolivia is post-GATT accession but those tariffs were set before GATT accession and unchanged between 1990-93. (d) The Czech Republic entered the GATT as a sovereign country in 1993. Its tariffs in 1992 were common to Slovakia with which it had a federation, which was a GATT-member. So it is possible that the tariffs for this country do not reflect a terms-of-trade motive. Our results by country in table 9 support this. Moreover, as we note in section 5.3, the pooled tariff results are robust to dropping the Czech Republic.

Table 2
Trade and Tariff Data Summary Statistics

	Trade Data			Tariff Data				
				Rate per 4-digit HS				Fraction of HS6 variation
	Number of Varieties*	Number of HS4 goods	Median # of Var per HS4	No. obs**	Mean	St. Dev.	Median	between HS4
Algeria	26466	1100	13	739	23.8	17.4	15.6	0.95
Belarus	24440	1172	12	703	12.4	7.8	10.0	0.94
Bolivia	18592	1064	9	647	9.8	0.8	10.0	0.63
China	63764	1217	33	1125	37.9	26.0	30.3	0.93
Czech	61781	1219	30	1075	9.5	17.6	5.1	0.87
Ecuador	22979	1101	11	753	9.8	5.5	10.6	0.91
Latvia	33790	1128	17	872	7.3	10.5	1.0	0.90
Lebanon	34187	1109	15	782	17.1	14.8	15.0	0.87
Lithuania	34825	1159	17	811	3.6	7.4	0.0	0.90
Oman	20482	1107	10	629	5.7	8.7	5.0	0.76
Paraguay	15430	1049	7	511	16.1	11.3	14.0	0.91
Russian	66731	1187	34	1029	10.7	11.0	5.7	0.95
Saudi Arabia	62525	1202	32	1036	12.1	2.6	12.0	0.93
Taiwan	38397	1215	19	891	9.7	8.5	7.5	0.90
Ukraine	37693	1128	18	730	7.4	7.6	5.0	0.95
Median	34187	1128	17	782	9.8	8.7	10.0	0.91

Notes: * Varieties are defined as 6-digit HS, exporting country pairs. **Number of observations for which elasticities and tariffs are available.

Table 3A
Inverse Export Supply Elasticity Statistics

Statistic	No. obs*	Median**			Mean		St. Deviation	
Sample	All	Low	Medium	High	All	W/out top decile	All	W/out top decile
Algeria	739	0.4	2.8	91	118	23	333	47
Belarus	703	0.3	1.5	61	85	15	257	36
Bolivia	647	0.3	2	91	102	23	283	49
China	1125	0.4	2.1	80	92	17	267	35
Czech Rep.	1075	0.3	1.4	26	63	7	233	18
Ecuador	753	0.3	1.5	56	76	13	243	30
Latvia	872	0.2	1.1	9	52	3	239	8
Lebanon	782	0.1	0.9	31	56	7	215	18
Lithuania	811	0.3	1.2	24	65	6	235	16
Oman	629	0.3	1.2	25	209	7	3536	21
Paraguay	511	0.4	3	153	132	67	315	169
Russia	1029	0.5	1.8	33	48	8	198	18
Saudi Arabia	1036	0.4	1.7	50	71	11	232	25
Taiwan	891	0.1	1.4	131	90	20	241	43
Ukraine	730	0.4	2.1	78	86	16	254	34
Median	782	0.3	1.6	54	85	13	243	30

Notes: *Number of observations for which elasticities and tariffs are available. The tariff availability did not bind except for Ukraine where it was not available for about 130 HS-4 goods for which elasticities were computed. ** The median over the "Low" sample corresponds to the median over the bottom tercile of inverse elasticities. Medium and High correspond to the 2nd and 3rd terciles.

Table 3B
Bootstrapped Statistics for Inverse Export Supply Elasticities

Sample	Low		Medium or High	
Statistic	Median	Confidence Interval*	Median	Confidence Interval*
Algeria	0.5	[0.10 , 0.8]	5.0	[2.0 , 81]
Belarus	0.3	[0.03 , 0.5]	3.0	[0.9 , 58]
Bolivia	0.4	[0.02 , 0.6]	4.2	[1.1 , 87]
China	0.6	[0.15 , 0.8]	5.0	[1.5 , 59]
Czech Rep.	0.3	[0.06 , 0.5]	3.0	[0.9 , 30]
Ecuador	0.4	[0.02 , 0.6]	3.3	[0.9 , 63]
Latvia	0.3	[0.02 , 0.4]	2.3	[0.7 , 6]
Lebanon	0.2	[0.01 , 0.3]	2.1	[0.6 , 29]
Lithuania	0.3	[0.03 , 0.5]	2.3	[0.7 , 28]
Oman	0.4	[0.04 , 0.6]	2.2	[0.6 , 35]
Paraguay	0.5	[0.03 , 0.8]	6.7	[1.9 , 98]
Russia	0.6	[0.12 , 0.7]	3.8	[1.3 , 42]
Saudi Arabia	0.5	[0.10 , 0.7]	4.1	[1.4 , 44]
Taiwan	0.3	[0.01 , 0.3]	3.0	[0.8 , 98]
Ukraine	0.6	[0.08 , 0.9]	4.5	[1.4 , 59]
Median	0.4	[0.04 , 0.6]	3.4	[1.1 , 49]

"Median" indicates the median of the 250 bootstrapped estimates for each inverse elasticity, ωig. The "low" column reports the median of that value in the bottom tercile of the sample for ωig in a country. The value in the medium or high column corresponds to the median in the rest of the sample.

* The lower bound of the confidence interval reported is the median lower bound over all those estimated in the relevant part of the sample. Similarly for the upper bound. The individual estimates for the 1-2α confidence interval are obtained via the bias-corrected percentile method (Efron, 1981) using α=0.1.

Table 4
Correlation of Inverse Export Supply Elasticities Across Countries

Dependent Variable:	Log Inverse Export Supply			
Statistic	Beta	St. Error	R2	Nobs
Algeria	0.80	(0.07)	0.13	739
Belarus	0.80	(0.07)	0.14	703
Bolivia	0.82	(0.09)	0.13	647
China	0.54	(0.06)	0.11	1125
Czech Rep.	0.61	(0.05)	0.12	1075
Ecuador	0.73	(0.08)	0.12	753
Latvia	0.57	(0.07)	0.09	872
Lebanon	0.71	(0.08)	0.11	782
Lithuania	0.70	(0.07)	0.13	811
Oman	0.39	(0.08)	0.04	629
Paraguay	0.94	(0.11)	0.14	511
Russia	0.53	(0.05)	0.11	1029
Saudi Arabia	0.48	(0.06)	0.08	1036
Taiwan	0.31	(0.08)	0.02	891
Ukraine	0.83	(0.07)	0.17	730
Median	0.70	(0.07)	0.12	782

* Univariate regression of log inverse export supply elasticities in each country on the average of the log inverse elasticities in that good for the remaining 14 countries.

Table 5
Inverse Elasticities by Product Type

	Differentiated	Reference Priced	Commodity
Median Inv Elasticity	2.38	0.70	0.45
Standard errors	(0.04)	(0.06)	(0.14)
p-value: differentiated vs. refer. or commod.		0.00	0.00
Mean Inv Elasticity	17.5	9.3	8.3
Standard errors	(0.71)	(0.70)	(1.23)
p-value: differentiated vs. refer. or commod.		0.00	0.00

Notes:

The number of observations for the median regression is 8734, less than the full sample since not all hs 4 can be uniquely matched to Rauch's classification.

The number for the mean regression is 7927 because we trim the top decile. The pattern of results with the top decile is similar but with higher values.

Table 6
Inverse Export Supply Elasticities, GDP, Remoteness and Import Shares

Dependent Variable:		Log Inverse Export Supply	
Log GDP	0.17	0.18	
	(0.04)	(0.03)	
Log Remoteness		0.40	
		(0.15)	
Share of World HS-4 Imports			7.19
			(1.48)
Observations	12343	12343	12343
R-square	0.26	0.26	0.25
R-square within	0.01	0.02	0.00

Notes: All regressions include 4-digit HS fixed effects (1201 categories). Robust standard errors in parentheses. In the log GDP regressions, standard errors are clustered by country. GDP is for 1996. Remoteness for country i is defined as 1/(ΣjGDPj/distance ij). The share of world imports is calculated in 2000.

Table 7
Tariffs and Market Power Across Goods (within countries): OLS & Tobit Estimates

Dependent Variable	Average Tariff at 4-digit HS (%)								
Fixed Effects	Country			Country & Industry					
Estimation Method	OLS	OLS	OLS	OLS	OLS	OLS	Tobit	OLS*	OLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Inverse Exp. Elast.	0.0003			0.0004					
	(0.0001)			(0.00004)					
Mid and High Inv Exp Elast		1.24			1.46			1.86	
		(0.25)			(0.24)			(0.31)	
Log(1/Export Elasticity)			0.12			0.17	0.17		
			(0.04)			(0.04)	(0.05)		
(Inv. Exp. Elast)*(1-med hi)								1.45	
								(0.31)	
(Inv. Exp. Elast)*med hi								0.0003	
								(0.0001)	
Mid Inv Exp Elast									1.56
									(0.28)
High Inv Exp Elast									1.37
									(0.28)
Algeria	23.8	23.0	23.6	24.6	23.6	24.3	24.3	23.1	23.6
	(0.64)	(0.65)	(0.64)	(0.95)	(0.96)	(0.95)	(0.93)	(0.97)	(0.96)
Belarus	12.3	11.5	12.2	12.6	11.6	12.5	12.4	11.3	11.7
	(0.29)	(0.33)	(0.29)	(0.76)	(0.78)	(0.76)	(0.94)	(0.79)	(0.78)
Bolivia	9.8	9.0	9.7	10.1	9.2	10.0	10.0	8.8	9.2
	(0.03)	(0.17)	(0.06)	(0.73)	(0.75)	(0.73)	(0.95)	(0.77)	(0.75)
China	37.8	37.0	37.7	38.2	37.2	38.0	37.9	36.6	37.2
	(0.77)	(0.79)	(0.77)	(0.98)	(1.01)	(0.99)	(0.89)	(1.03)	(1.01)
Czech Rep.	9.5	8.7	9.4	9.7	8.7	9.6	8.8	8.3	8.7
	(0.53)	(0.53)	(0.53)	(0.85)	(0.86)	(0.85)	(0.89)	(0.87)	(0.86)
Ecuador	9.8	9.0	9.7	10.3	9.4	10.2	10.1	9.0	9.4
	(0.19)	(0.26)	(0.20)	(0.73)	(0.74)	(0.73)	(0.93)	(0.76)	(0.74)
Latvia	7.3	6.4	7.2	7.3	6.3	7.2	6.9	6.0	6.3
	(0.35)	(0.40)	(0.35)	(0.76)	(0.78)	(0.76)	(0.91)	(0.79)	(0.78)
Lebanon	17.1	16.2	17.0	17.1	16.1	17.0	17.0	15.9	16.1
	(0.53)	(0.56)	(0.53)	(0.84)	(0.86)	(0.84)	(0.92)	(0.86)	(0.86)
Lithuania	3.6	2.8	3.6	3.6	2.6	3.5	-6.0	2.3	2.6
	(0.26)	(0.31)	(0.26)	(0.74)	(0.76)	(0.74)	(0.98)	(0.77)	(0.76)
Oman	5.6	4.9	5.6	5.7	4.8	5.6	4.9	4.4	4.8
	(0.34)	(0.37)	(0.34)	(0.77)	(0.79)	(0.77)	(0.94)	(0.79)	(0.79)
Paraguay	16.0	15.3	15.9	16.3	15.4	16.1	15.9	14.9	15.4
	(0.49)	(0.52)	(0.50)	(0.84)	(0.85)	(0.84)	(0.99)	(0.86)	(0.85)
Russia	10.6	9.8	10.5	10.8	9.9	10.7	10.0	9.4	9.9
	(0.34)	(0.38)	(0.34)	(0.77)	(0.79)	(0.77)	(0.89)	(0.82)	(0.79)
Saudi Arabia	12.1	11.3	12.0	12.4	11.4	12.2	12.1	10.9	11.4
	(0.08)	(0.18)	(0.09)	(0.71)	(0.74)	(0.72)	(0.89)	(0.76)	(0.74)
Taiwan	9.7	8.9	9.6	10.3	9.3	10.1	9.7	9.0	9.3
	(0.28)	(0.33)	(0.28)	(0.74)	(0.76)	(0.75)	(0.91)	(0.77)	(0.76)
Ukraine	7.4	6.6	7.2	8.1	7.1	7.9	6.8	6.6	7.1
	(0.28)	(0.33)	(0.29)	(0.74)	(0.76)	(0.74)	(0.93)	(0.78)	(0.76)
Observations	12333	12333	12333	12333	12333	12333	12333	12333	12333
Number of parameters	16	16	16	36	35	36	35	38	36
Adj. R2	0.61	0.61	0.61	0.66	0.66	0.66	.	.	0.66

Notes: Standard errors in parenthesis (all heteroskedasticity robust except Tobit). Industry dummies defined by section according to Harmonized Standard tariff schedule. * Optimal threshold regression based on minimum RSS found using a grid search over 50 points of the distribution of inverse exp. elast (from 1st to 99th percentile in intervals of 2). Optimal threshold is 53rd percentile. Accordingly, med hi=1 above the 53rd percentile and 0 otherwise. Hansen (2000) shows that the dependence of the parameters on the threshold estimate is not of "first-order" asymptotic importance, so inference on them can be done as if the threshold estimate were the true value.

Table 8
Tariffs and Market Power Across Goods (within countries): IV Estimates

Panel A: Full sample

Dependent Variable	Average Tariff at 4-digit HS (%)								
Fixed Effects	Country			Country & Industry			Industry by Country		
Estimation method	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Inverse Exp. Elast.	0.040			0.089			0.075		
	(0.027)			(0.055)			(0.028)		
Mid and High Inv Exp Elast		3.96			8.88			9.07	
		(0.76)			(1.18)			(1.08)	
Log(1/Export Elasticity)			0.75			1.71			1.73
			(0.15)			(0.23)			(0.21)
Observations	12258	12258	12258	12258	12258	12258	12258	12258	12258
no. of parameters	16	16	16	35	35	35	284	282	283
1st stage F	5	1649	1334	2	653	517	3	691	544

Panel B: Sub sample (exc. Bolivia, S. Arabia, Oman)

Dependent Variable	Average Tariff at 4-digit HS (%)								
Fixed Effects	Country			Country & Industry			Industry by Country		
Estimation Method	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Inverse Exp. Elast.	0.031			0.065			0.067		
	(0.006)			(0.012)			(0.011)		
Mid and High Inv Exp Elast		5.17			10.64			10.99	
		(0.89)			(1.37)			(1.27)	
Log(1/Export Elasticity)			0.97			2.04			2.11
			(0.17)			(0.27)			(0.25)
Observations	9952	9952	9952	9952	9952	9952	9952	9952	9952
no. of parameters	13	13	13	32	32	32	227	226	227
1st stage F	129	1448	1187	48	580	456	50	611	477

Notes: Standard errors in parenthesis (heteroskedasticity robust).Industry dummies defined by section according to the Harmonized Standard tariff schedule.

Table 9
Tariffs and Market Power Across Goods by country: IV Estimates

Panel A: Semilog

Dependent Variable	Average Tariff at 4-digit HS (%)																	
Fixed Effects	Ind. by Country	Ind. by Country	Ind.	Ind.	Ind.		Ind.	Ind.	Ind.	Ind.	Ind.		Ind.	Ind.	Ind.		Ind.		
Estimation Method	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV Tobit	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV Tobit	IV GMM	IV GMM	IV GMM	IV Tobit	IV GMM		
Sample	All	Exc. Bol., Om., Sau.	China	Russia	Taiwan	Saudi Arabia	Ukraine	Czech	Algeria	Belarus	Ecuador	Oman	Paraguay	Lithuania	Lebanon	Bolivia	Latvia	Mean	Median
Log(1/Export Elasticity)	1.73	2.11	7.60	2.42	1.98	1.75	0.71	0.16	5.40	2.28	1.55	0.60	2.44	0.83	2.41	0.79	1.41	2.15	1.75
	(0.21)	(0.25)	(1.78)	(0.61)	(0.78)	(2.55)	(0.26)	(0.24)	(0.97)	(0.52)	(0.34)	(0.18)	(0.67)	(0.27)	(0.54)	(0.36)	(0.60)		
Observations	12258	9952	1089	1021	841	1031	685	1000	739	703	753	628	510	768	754	647	868		
no. of parameters	283	227	20	18	20	.	18	20	18	18	19	.	17	19	20	.	19		
1st stage F	544	477	45	44	12	7.6*	48	60	60	40	45	3.7*	33	35	52	8.96*	37		
Mean Tariff (%)	13.4	14.2	38.2	10.3	8.9	12.2	5.8	5.5	23.8	12.4	9.8	5.7	16.0	2.3	15.0	9.8	7.0		
Elasticity (at mean)	0.13	0.15	0.20	0.23	0.22	0.14	0.12	0.03	0.23	0.18	0.16	0.11	0.15	0.36	0.16	0.08	0.20	0.17	0.16

Panel B: Dummy

Dependent Variable	Average Tariff at 4-digit HS (%)																	
Fixed Effects	Ind. by Country	Ind. by Country	Ind.	Ind.	Ind.		Ind.	Ind.	Ind.	Ind.	Ind.		Ind.	Ind.	Ind.		Ind.		
Estimation Method	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV Tobit	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV Tobit	IV GMM	IV GMM	IV GMM	IV Tobit	IV GMM		
Sample	All	Exc. Bol., Om., Sau.	China	Russia	Taiwan	Saudi Arabia	Ukraine	Czech	Algeria	Belarus	Ecuador	Oman	Paraguay	Lithuania	Lebanon	Bolivia	Latvia	Mean	Median
Mid and High Inv Exp Elast	9.07	10.99	35.3	11.3	11.2	7.3	3.6	0.8	28.7	12.7	8.6	2.5	19.0	3.7	16.3	5.6	6.3	12	9
	(1.08)	(1.27)	(7.73)	(2.85)	(3.88)	(10.84)	(1.32)	(1.17)	(4.78)	(2.90)	(1.81)	(0.57)	(5.86)	(1.19)	(3.65)	(2.44)	(2.61)		
Observations	12258	9952	1089	1021	841	1031	685	1000	739	703	753	628	510	768	754	647	868		
no. of parameters	282	226	20	18	20	.	18	20	18	18	19	.	17	19	19	.	19		
1st stage F	691	611	62	54	17	10.7*	64	77	84	47	56	5.2*	22	52	46	8.5*	57		
Mean Tariff (%)	13.4	14.2	38.2	10.3	8.9	12.2	5.8	5.5	23.8	12.4	9.8	5.7	16.0	2.3	15.0	9.8	7.0		
mid-hi/mean (%)	68	77	92	110	126	60	61	14	121	103	87	44	119	163	109	57	90	90	92

Notes: Standard errors in parenthesis (heteroskedasticity robust). Industry dummies defined by section according to the Harmonized Standard tariff schedule. Extreme tariff outliers were dropped. This affects the sample of only 5 countries and even then only slightly since the criteria only drops 3-7% of their observations (those with tariff values more than 3 times the interquartile range above the 75th percentile or below the 25th). Bolivia, Oman and Saudi Arabia are estimated via an IV Tobit procedure to account for the fact that a large fraction of their observations are censored from above and/or below. Given the lack of variation in their tariff within industries their estimation does not include industry dummies.

* z-stat of the instrument in the first stage of IV Tobit.

Table 10
Market Power vs. Tariff Revenue or Lobbying as a Source of Protection

Dependent Variable	Average Tariff at 4-digit HS (%)					
Fixed Effects	Industry by Country					
Estimation Method	IV GMM					
Sample	Pooled (all)		Pooled (all)		Pooled (7)	
Theory	*Market Power*		*Market Power and Tariff Revenue*		*Market Power and Lobbying*	
Mid and High Inv Exp Elast	9.07		9.04		10.20	
	(1.08)		(1.24)		(1.79)	
Mid and High Inv Imp Elast			-0.20			
			(2.08)			
Mid and hi Inv. Imp. Pen/Imp. elast					6.28	
					(1.97)	
Log(1/Export Elasticity)		1.73		1.81		1.94
		(0.21)		(0.23)		(0.38)
Log(1/Import Elasticity)				-0.90		
				(0.81)		
Log(Inv. Imp. Pen/Imp. elas.)						1.59
						(0.55)
Observations	12258	12258	12258	12258	5178	5178
no. of parameters	282	283	283	284	132	133
1st stage F (Market power)	691	544	370	312	171	129
1st stage F (other)	na	na	102	144	131	188

Notes: Standard errors in parenthesis (heteroskedasticity robust).Industry dummies defined by section according to the Harmonized Standard tariff schedule. The countries with available data for the lobbying specifications are Bolivia, China, Ecuador, Latvia Lithuania, Taiwan and Ukraine. This data is not available for mining and agricultural products.

Table 11
Market Power and Lobbying: IV Estimates by Country

Panel A: Semilog

Dependent Variable	Average Tariff at 4-digit HS (%)									
Fixed Effects	Ind. by Country	Ind.	Ind.	Ind.	Ind.	Ind.		Ind.		
Estimation Method	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV Tobit	IV GMM		
Sample	Pooled (7)	China	Taiwan	Ukraine	Ecuador	Lithuania	Bolivia	Latvia	Mean	Median
Log(1/Export Elasticity)	1.94	4.69	2.39	0.91	1.81	0.84	0.97	1.52	1.87	1.5
	(0.38)	(2.12)	(1.32)	(0.25)	(0.45)	(0.27)	(0.16)	(0.67)		
Log(Inv. Imp. Pen/Imp. elas.)	1.59	6.21	0.43	0.97	0.27	1.64	0.21	1.89	1.66	1.0
	(0.55)	(4.31)	(1.18)	(0.75)	(0.57)	(0.40)	(0.19)	(1.33)		
Observations	5178	861	780	616	712	706	618	788		
no. of parameters	133	21	20	19	20	20	.	20		
1st stage F: log(1/exp. El)	129	38	6	25	24	18	9	18		
1st st. F: log(Imp. pen/imp. el)	188	37	32	7	47.2	28.5	18.4	24.9		
Mean Tariff (%)	12.8	37.0	8.8	5.7	10.0	2.4	9.8	6.9		
Elasticity (at mean)	0.15	0.13	0.27	0.16	0.18	0.35	0.10	0.22	0.20	0.18

Panel B: Dummy

Dependent Variable	Average Tariff at 4-digit HS (%)									
Fixed Effects	Ind. by Country	Ind.	Ind.	Ind.	Ind.	Ind.		Ind.		
Estimation Method	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV GMM	IV Tobit	IV GMM		
Sample	Pooled (7)	China	Taiwan	Ukraine	Ecuador	Lithuania	Bolivia	Latvia	Mean	Median
Mid and High Inv Exp Elast	10.2	22.9	13.3	4.2	10.3	3.4	8.0	6.7	9.83	8.0
	(1.79)	(9.18)	(4.61)	(1.17)	(2.87)	(1.43)	(0.92)	(2.86)		
Mid and hi Inv. Imp. Pen/Imp. elast	6.28	16.1	0.9	1.4	1.5	6.4	2.5	6.4	5.01	2.5
	(1.97)	(10.07)	(3.52)	(1.84)	(3.35)	(1.84)	(1.36)	(3.54)		
Observations	5178	861	780	616	712	706	618	788		
no. of parameters	132	21	20	19	20	20	.	20		
1st stage F: log(1/exp. El)	171	48	10	36	27	24	9	29		
1st st. F: log(Imp. pen/imp. el)	131	37	18	11	23.9	14.0	17.8	19.7		
Mean Tariff (%)	12.8	37.0	8.8	5.7	10.0	2.4	9.8	6.9		
Mid-hi Inv Exp Elast /mean (%)	80	62	151	73	103	140	82	97	101	97

Notes: Standard errors in parenthesis (heteroskedasticity robust). Constant and industry dummies included but not reported. Industry dummies defined by section according to the Harmonized Standard tariff schedule. Observations with extreme outliers in terms of tariffs were dropped. This affects the sample for Taiwan, Lithuania and Ukraine but only slightly since the criteria drops less than 36 observations in any of these countries (those with tariff values more than 3 times the interquartile range above the 75th percentile or below the 25th). Bolivia is estimated via an IV Tobit procedure to account for the fact that a large fraction of their observations are censored from above and/or below. Given the lack of variation in its tariffs within industries the estimation for Bolivia does not include industry dummies.

* z-stat of the relevant instrument in first stage.

Table 12
Economic and Statistical Importance of Market Power in Tariff Setting: Summary Measures

		Specification				
		Market Power			*Market Power and Lobbying*	
Variable	Measure	Pooled (Table 8A)	Subsample (Table 8B)	Typical Country (Table 9)	Pooled (Table 10)	Typical Country (Table 11)
Log(1/Export Elasticity)	β	1.7 pp	2.1 pp	1.8 pp	1.9 pp	1.5 pp
	Elasticity (β/mean tariff)	0.13	0.15	0.16	0.15	0.18
	β*s.d.	5 pp	6 pp	5 pp	5 pp	4 pp
	Elasticity relative to PE (β/γ)	.	.	.	1.2	0.9
	Impact relative to PE (β*s.d.(lnω)/γ*s.d.(lnz/σ))	.	.	.	1.6	1.5
Mid and High Inv Exp Elast	β	9 pp	11 pp	9 pp	10 pp	8 pp
	β/mean tariff (%)	68%	77%	92%	80%	97%
	Impact relative to PE (β/γ)	.	.	.	1.6	3.1

Note: β and γ correspond to the coefficients on market power and the lobbying variable respectively. "pp" stands for percentage points.

Table 13
Market Power and Lobbying as a Source of Protection in the US

Panel A: Non-Tariff Barriers

Theory	*Market Power*				*Market Power and Lobbying*			
Fixed Effects	Industry				Industry			
Estimation Method	IV Tobit				IV Tobit**			
Dependent Variable	Coverage Ratio (HS-4)*		Advalorem equiv. (HS-4, %)		Coverage Ratio (HS-4)		Advalorem equiv. (HS-4, %)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Mid and High Inv Exp Elast	0.90		38.8		4.93		70.8	
	(0.31)		(15.73)		(1.52)		(21.99)	
Mid and hi Inv. Imp. Pen/Imp. elast					-0.08		3.99	
					(0.86)		(13.14)	
Log(1/Export Elasticity)		0.22		9.71		1.16		16.0
		(0.08)		(4.00)		(0.39)		(5.47)
Log(Inv. Imp. Pen/Imp. elas.)						0.19		4.74
						(0.34)		(4.94)
Observations***	804	804	804	804	708	708	708	708
no. of parameters	17	17	17	17	17	17	17	17
1st stage z-stat (Market power)	7.1	6.6	7.1	6.6	6.2	5.3	6.2	5.3
1st stage z-stat (other)	na	na	na	na	10.1	11.4	10.1	11.4

Panel B: Tariff Barriers

Theory	*Market Power*				*Market Power and Lobbying*			
Fixed Effects	Industry				Industry			
Estimation Method	IV Tobit				IV Tobit**			
Dependent Variable	Non-WTO (HS-4, %)		WTO (HS-4, %)		Non-WTO (HS-4, %)		WTO (HS-4, %)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Mid and High Inv Exp Elast	21.2		1.52		26.9		1.89	
	(5.53)		(1.18)		(8.05)		(1.58)	
Mid and hi Inv. Imp. Pen/Imp. elast					10.8		-0.63	
					(4.91)		(0.96)	
Log(1/Export Elasticity)		5.07		0.36		5.58		0.45
		(1.36)		(0.28)		(1.86)		(0.38)
Log(Inv. Imp. Pen/Imp. elas.)						4.76		-0.18
						(1.69)		(0.34)
Observations***	870	870	869	869	775	775	774	774
no. of parameters	20	20	20	20	21	21	21	21
1st stage z-stat (Market power)	7.3	7.1	7.3	7.1	6.0	5.3	6.0	5.3
1st stage z-stat (other)	na	na	na	na	10.0	11.6	10.0	11.6
Mean	30.6	30.6	3.4	3.4	33.0	33.0	3.7	3.7
Mid-hi Inv Exp Elast /mean (%)	69		45		81		51	
Elasticity (at mean)		0.17		0.11		0.17		0.12

Notes: Standard errors in parenthesis. Industry dummies defined by section according to the Harmonized Standard tariff schedule.

* Coverage ratio is defined as the fraction of HS-6 digit tariff lines in a given HS-4 category that had an NTB. Since it varies between zero and 1 we use a two-limit IV Tobit. For the remaining variables we use a lower limit Tobit that accounts for censoring at zero. There is a lower share of censored observations in Panel B, and we confirmed that these results are very similar if we use IV-GMM instead.

** We employ the Newey two-step estimator in the specifications with more than one endogenous variables since it is well known that in these cases the maximum likelihood estimator has difficulty in converging.

*** The difference in the number of observations across specifications is due to missing production data for mining and agricultural products. The difference between tariff and non-tariff barriers is due to the lack of variation of NTBs within certain industries, which must therefore be dropped. The tariff results in Panel B based on a comparable sample to the NTB are identical.

Appendix I: Optimal tariff with CES utility for foreign varieties

We employ a utility function $u = u(M^{\mu}D^{\alpha})$ for the non-numeraire goods where M is the subutility function for imported goods and D is a composite domestic good. This is consistent with the empirical approach we employ to estimate the elasticities. We rule out income effects by using the quasilinear structure outlined in the text. Moreover, we take M as a Cobb-Douglas aggregate over imported goods $g \in G_m$, so $M = \prod_g M_g^{\phi_g}$. Each M_g is in turn composed of varieties that are aggregated via a CES utility, $M_g = (\Sigma_v d_{gv}^{1-\kappa_g} m_{gv}^{\kappa_g})^{1/\kappa_g}$ where $d_{gv} > 0$ is a taste or quality parameter for gv and $1/(1-\kappa_g)$ represents the constant elasticity of substitution, σ_g. Finally, we assume that the subutility represented by u gives rise to constant expenditure. That is $u = \ln(M^{\mu}D^{\alpha})$. Thus we can rewrite the utility as

$$U = c_0 + \mu \sum_{g \in G_m} \phi_g \ln[(\Sigma_v d_{gv}^{1-\kappa_g} m_{gv}^{\kappa_g})^{1/\kappa_g}] + \alpha \ln D \; ; \; 0 < \kappa_g < 1$$

This structure implies separability across imported goods but not its varieties. We could provide a similar treatment to the domestic composite good but since it is log separable this is not necessary in order to determine the optimal tariffs for each imported good. As shown in the standard model in the text a demand structure with quasilinear utility and separability over goods g implies that tariffs in any good g only impact welfare through the consumption of that good and the tariff revenue it generates. Thus we can focus on deriving the optimal tariff separately for each good $g \in G_m$. Furthermore, we assume that $\tau_{gv} = \tau_g$ for all v in a given good.

We now show that the optimal tariff for each of these goods is given by the inverse export supply elasticity as we estimate it. In the estimation we must assume the elasticities are constant and identical within goods across imported varieties. Therefore we impose those conditions and write the export supply for each variety of good g as

$$m_{gv}^{*} = a_{gv} p_{gv}^{*1/\omega_g} \text{ all } g, v$$

Note that different exporters may have different a_{gv} and this variation, along with the differences in demand taste generate variation in prices across exporters that is important in identifying the elasticities empirically (we also use the time variation). The demand for each variety is obtained by minimizing expenditure subject to obtaining a given level M_g. It is given by

$$\begin{aligned} m_{gv} &= d_{gv}\left(\frac{p_g}{p_{gv}}\right)^{\sigma_g} M_g \\ &= \mu\phi_g d_{gv} p_{gv}^{-\sigma_g} p_g^{\sigma_g - 1} \end{aligned}$$

where p_g represents the standard price index, $\left(\sum_v d_{gv} p_{gv}{}^{1-\sigma_g}\right)^{1/(1-\sigma_g)}$. The second line uses the fact that $u()$ has a constant expenditure share and that the maximization of that utility will yield $p_g M_g = \mu\phi_g$. Using this and the export supply equation we obtain the market clearing price obtained by foreign exporters, $p_{gv}^{*} = p_{gv}/(1+\tau_{gv})$.

$$\mu\phi_g d_{gv}(p_{gv}^{*}(1+\tau_g))^{-\sigma_g} p_g^{\sigma_g - 1} = a_{gv} p_{gv}^{*1/\omega_g}$$

$$p_{gv}^{*} = [\varphi_{gv}(1+\tau_g)^{-\sigma_g} p_g^{\sigma_g - 1}]^{\frac{\omega_g}{1+\omega_g\sigma_g}}$$

where $\varphi_{gv} \equiv \frac{\mu\phi_g d_{gv}}{a_{gv}}$ depends on variety specific characteristics.

The key insight to showing that the optimal tariff equals ω_g is to note that the tariff does not affect the relative demand of varieties in any given good. Therefore the only distortion that it addresses is the terms of trade externality. There are three assumptions that are required for this result. First, consumption and export supply elasticities within a good are constant. Second, they are identical across varieties or exporters of that good. Third, tariffs are equal across exporters of the same good. To see this we can simply use the expressions for m_{gv}, and p_{gv}^{*} to obtain the relative demand across any two varieties v, k of a given good as

$$\frac{m_{gv}}{m_{gk}} = \frac{d_{gv}}{d_{gk}}\left(\frac{\varphi_{gk}}{\varphi_{gv}}\right)^{\frac{\omega_g\sigma_g}{1+\omega_g\sigma_g}}$$

To obtain individual prices, p_{gv}^{*}, as a function of tariffs we first solve for the aggregate price index of each good and then replace it in the expression for p_{gv}^{*}. To do so we first note that $p_g = (1+\tau_g)\left(\sum_v d_{gv} p_{gv}^{*\,1-\sigma_g}\right)^{1/(1-\sigma_g)}$ and then aggregate the individual prices from the market clearing conditions to obtain an expression similar to this one, which can be solved to obtain

$$p_g(\tau_g) = (1+\tau_g)^{\frac{1}{1+\omega_g}}\Phi$$

where $\Phi \equiv \left(\sum_v d_{gv}[\varphi_{gv}]^{\frac{\omega_g(1-\sigma_g)}{(1+\omega_g\sigma_g)}}\right)^{(1+\sigma_g\omega_g)/(1+\omega_g)(1-\sigma_g)}$. We can verify that if $\omega_g = 0$ there is complete pass-through from tariffs to the aggregate price of g and as ω_g increases the effect of tariffs to domestic prices is attenuated. Replacing this in the market clearing condition for each variety we obtain $p_{gv}^{*}(\tau_g)$.

$$p_{gv}^{*} = (1+\tau_g)^{-\frac{\omega_g}{1+\omega_g}}[\varphi_{gv}\Phi^{\sigma_g-1}]^{\frac{\omega_g}{(1+\omega_g\sigma_g)}}$$

The pass-through of the tariff *factor* to domestic prices is simply $d\ln p_{gv}/d\ln(1+\tau_g) = 1 + d\ln p_{gv}^{*}/d\ln(1+\tau_g) = 1/(1+\omega_g)$, where the last equality follow from differentiating the last equation. Had we explicitly included the exchange rate, its pass-through would be exactly symmetric to the tariff factor. The elasticity of the exporter price wrt the tariff itself is $d\ln p_{gv}^{*}/d\ln\tau_g = -\frac{\omega_g}{1+\omega_g}\frac{\tau_g}{1+\tau_g}$.

The equilibrium imports as a function of tariffs is then $m_{gv} = m_{gv}^{*}$, for each variety of g given by

$$m_{gv}^{*}(\tau_g) = (1+\tau_g)^{-\frac{1}{1+\omega_g}}\Gamma_{gv}$$

$\Gamma_{gv} \equiv a_{gv}[\varphi_{gv}\Phi^{\sigma_g-1}]^{\frac{1}{1+\omega_g\sigma_g}}$

The government will then choose τ_g for each good to maximize the following social welfare expression

$$\max_{\tau_g} W = 1 + \sum_g W_g + \pi_d + s_d$$

where for each good g we have that W_g is given by $\tau_g \Sigma_v p^*_{gv}(\tau_g) m_{gv}(\tau_g) + \mu\phi_g \ln[(\Sigma_{gv} d_{gv}^{1-\kappa_g} m_{gv}(\tau_g)^{\kappa_g})^{1/\kappa_g}] - \Sigma_{gv} p_{gv}(\tau_g) m_{gv}(\tau_g)$. The foc for each good g can then be derived and simplified to obtain

$$\Sigma_v \left(\tau_g p^*_{gv} \frac{dm_{gv}}{d\tau_g} - m_{gv} \frac{dp^*_{gv}}{d\tau_g} \right) = 0$$

Therefore we obtain an expression similar to the one in the text under the standard model. However, now it is defined over the sum of the varieties. To see that the elasticity we estimate is exactly the solution we can rewrite the expression above in terms of the elasticities of m_{gv} and p^*_{gv} wrt τ_g. Using the equilibrium level of imports and prices derived it is simple to see that these are constant across varieties in a good and thus we obtain the inverse elasticity solution.

$$\Sigma_v p^*_{gv} m_{gv} \left(\frac{dm_{gv}}{d\tau_g} \frac{\tau_g}{m_{gv}} - \frac{dp^*_{gv}}{d\tau_g} \frac{\tau_g}{p^*_{gv}} \frac{1}{\tau_g} \right) = 0$$

$$\Sigma_v p^*_{gv} m_{gv} \left(-\frac{1}{1+\omega_g} \frac{\tau_g}{1+\tau_g} - (-\frac{\omega_g}{1+\omega_g} \frac{\tau_g}{1+\tau_g}) \frac{1}{\tau_g} \right) = 0$$

This implies that $\tau_g = \omega_g$ for all g.

Appendix II: Testing the independence of errors assumption

The estimation of elasticities relies on the independence of errors across relative demand and supply shocks, $E_t\left(\varepsilon_{ivgt}^{k_{ig}}\delta_{ivgt}^{k_{ig}}\right)=0$. As we describe in that section the double differencing eliminates many possible reasons why such shocks would be correlated. However, there is one potential case that would lead to such a correlation. In this section, we test whether it affects our estimates.

New sub-varieties of a particular variety (e.g. new HS-10 products within a given HS 6-digit of a given exporter) may appear as a demand shock (ε_{ivgt}). Moreover, in some models of trade, productivity shocks (which are one component of the error in the supply equation, δ_{ivgt}) will cause the number of exported varieties to rise. If these shocks occurred in all varieties (i.e. exporters) of *g* the independence condition would still be satisfied. However, if the shocks differ across varieties it would not. Naturally, there are other supply shocks, which may be far more important determinants of relative costs in the real world. In the case of our data we think that exchange rate shocks are one source of large variation at the yearly frequency that we exploit. So even if the case above is a theoretical possibility, it is not necessarily important in the data. . This is what we now test.

The concern can be thought of in terms of the level of aggregation. If we could observe all sub-varieties (i.e., sub-varieties within 6-digit varieties), then there would be no problem. Our data already breaks up the typical countries' imports into 30,000 varieties, so the scope for sub-varieties is likely to be limited because they are not likely to be that differentiated from existing ones. However, other 6-digit goods are more differentiated. For example, sector 62160 (gloves and mittens) contains at least two sub-varieties: gloves and mittens. Unfortunately, we cannot observe 6-digit sub-varieties for most of the countries in our sample because the data is not available. However, in the case of the U.S. we do have more detailed HS 10-digit data that allows us to directly evaluate how large is the bias in our point estimates from focusing on 6-digit vs. 10-digit data.

Our strategy for testing the validity of the error independence assumption is the following. We classify each U.S. 6-digit category by the role that sub-varieties play in its growth, and estimate 4-digit supply elasticities using *only* the sub-sample of 6-digit varieties in which changes in sub-varieties were negligible. We then compare these estimates, where we expect the identifying assumption to hold with the estimates obtained using *all* varieties. If the elasticity estimates are similar we can conclude that our identifying assumption also holds in the full sample. Thus we exploit the fact that we have over-identifying conditions for each elasticity (i.e. more varieties than typically needed for the estimation) and use the HS-10 digit data to identify the part of the sample where the identifying assumption is more likely to hold.

We measure the importance of the extensive margin for each 6-digit variety similarly to Feenstra (1994) and Hummels and Klenow (2005):

$$\text{(A1)}\quad \text{Measure of extensive margin}_{\text{vt}} = \left|\ln\left(\frac{m_{vt}\left(I_v\right)}{m_{vt}}\right)-\ln\left(\frac{m_{vt-1}\left(I_v\right)}{m_{vt-1}}\right)\right|,$$

where m_{vt} is the total imports of the U.S. of 6-digit variety *v* at time *t*, and I_v is the set of 10-digit sub-varieties of *v* that were imported in both periods *t and t-1*. Thus, $m_{vt}(I_v)$ is the value in period *t* of all sub-varieties of variety *v* that were common across periods. In particular, this measures the absolute change in the share of common sub-varieties between periods *t* and *t-1*. If the change in this share is zero, then there was no net change in the importance of sub-varieties within the variety. The larger the measure, the more important the extensive margin is.

Table A1 reports the histogram of the per-year average measure in (A1) across all varieties. In 80 percent of the HS-6 varieties the creation or destruction of sub-varieties moved the value of its sales by less than 2 percent. For around 90 percent of the varieties it is less than 9 percent. This indicates that for the vast majority of cases there are few changes in the importance of sub-varieties. Of course, the fact that movements in sub-varieties are unimportant for most varieties does not, by itself, prove that sub-varieties do not have an effect on our market power estimates. We therefore turn to estimating this impact.

Table A2 presents regressions that compare the 4-digit inverse supply elasticities estimated with the full sample of 6-digit varieties and the elasticities estimated using the restricted sample of 6-digit varieties where the extensive margin has little or no importance (and thus is less subject to the referee's criticism). We employ a cut-off of 5%, that is we eliminate any HS-6 where the extensive margin growth was more than 5%. The results are not sensitive to the exact cut-off.

To minimize the impact of influential outliers, we do the regression in logs. If sub-variety growth lead to a violation of our identifying assumption then it would generate biased estimates. This would imply that the elasticities based on the restricted sample would be systematically different from the estimates based on the full sample.

Column 1 shows that in the OLS regression the coefficient is 0.9 and the R^2 around 80 percent. This point estimate is likely to be attenuated because we do not adjust for measurement error in our independent variable. We easily confirm this is the case by running the reverse regression (i.e. placing the estimate that uses the restricted sample on the left-hand side) and finding the same coefficient. We correct for this measurement error bias in the second column by instrumenting for the log of the inverse export elasticity with the average log elasticity of the same 4-digit HS good obtained from a sample of 5 developed countries (Canada, France, Germany, Japan and the United Kingdom). This yields a point estimate of 1.0. This is strong evidence against the hypothesis that sub-variety growth is biasing our elasticity estimates.

Table A1

US Measure of Sub-variety per variety
(as defined in eq. (A1))

Percentile	Centile
50	0.00
60	0.00
70	0.00
80	0.02
90	0.09
95	0.19
99	1.72

Number of 6-digit varieties: 112616

Table A2

Dependent variable : lninv (US Sample)

Restriction	5 percent	5 percent
Type	OLS	IV
lninv (Restricted Sample)	0.9	1.0
	(0.015)	(0.034)
Observations	1038	1015
R-squared	0.78	
F 1st stage		269
Varieties Restricted / Full*	0.76	0.76

Standard errors in parentheses. *Represents the number of observations used to calculate the elasticities in the full relative to the restricted sample.