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Optimal Tariffs and Market Power: The Evidence

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We find that prior to World Trade Organization membership, countries set import tariffs 9 percentage points higher on inelastically supplied imports relative to those supplied elastically. The magnitude of this effect is similar to the size of average tariffs in these countries, and market power explains more of the tariff variation than a commonly used political economy variable. Moreover, US trade restrictions not covered by the WTO are significantly higher on goods where the United States has more market power. We find strong evidence that these importers have market power and use it in setting noncooperative trade policy. (*JEL* F12, F13)

The idea that countries set tariffs in response to their market power in international markets is a controversial result in international economics. For example, Kyle Bagwell and Robert W. Staiger (1999) argue that it provides the underlying motive for the world trading system, while Paul R. Krugman and Maurice Obstfeld (1997, 226) argue that "the terms-of-trade argument is of little practical importance." 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 foreign export supply elasticities?" The answer is none.[1]

The theory that a country with market power in trade might gain from protection has a long history but its main insight can be summarized as follows.[2] A tariff creates consumption and production distortions, but it also creates a terms-of-trade gain if the foreign supply is inelastic, i.e., if the importer has market power. Thus, in the absence of constraints such as trade agreements,

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[1] There is some evidence that changes in trade policy affect the prices of the goods that countries import (c.f. Mordechai E. Kreinin 1961; Won Chang and Alan L. Winters 2002; Robert C. Feenstra 1989). 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. Since our initial working paper, however, independent research by Bagwell and Staiger (2006) provides evidence for their theory by showing that WTO accession leads to greater tariff reductions in products with higher initial import volumes.

[2] Seminal contributions on this issue extend back to Robert Torrens (1833) and include John S. Mill (1844), Francis Y. Edgeworth (1894), Tibor Scitovsky (1942), and Harry G. Johnson (1954). Douglas A. Irwin (1996) carefully discusses the history of thought on optimal tariffs.

the theory predicts a positive relationship between a country's tariffs and its market power. The original derivation of such a relationship by Charles F. Bickerdike (1907) focused on the "optimal tariff" set by a welfare maximizing government. But the positive relationship between tariffs and market power also arises in more general settings that do not require welfare maximization, as we discuss in the theory section.

Our objective in this paper is to quantify the importance of the market power (or terms-of-trade) motive in trade policy. In doing so, we make three contributions. First, we estimate elasticities of export supply faced by 15 importer countries at a highly disaggregated level. Second, we use these elasticities to provide evidence that prior to constraints imposed by the World Trade Organization (WTO), these countries systematically set higher import tariffs on goods in which they have market power. Finally, we estimate similar elasticities for the United States and find that its trade restrictions that are not constrained by the WTO are significantly higher in goods where the United States has more market power. The results are robust to the inclusion of political economy variables and a variety of model specifications. The effect is statistically and economically significant relative to both 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.

We rely on the methodology of Feenstra (1994) and Broda and Weinstein (2006) to estimate the export supply elasticities at the four-digit Harmonized System (HS) level over the period 1994–2003. Our sample consists of the 15 countries for which we could obtain tariff data for a large fraction of products prior to constraints imposed by WTO membership. We find that the inverse export supply elasticity faced by an importer is between one and three for the typical four-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 indicate that, on average, they have more market power than small countries. 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.

Using these elasticities, we then estimate that, prior to WTO constraints, these countries set higher tariffs on products where they have more market power. This effect is present both when we compare median tariff rates across countries and when we compare actual tariff rates across HS four-digit goods within countries and industries. The impact of market power on tariffs is robust to many different specifications. The effect is present 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 countries and significant in 13 of the 15 countries studied. 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., Pinelopi K. Goldberg and Giovanni 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 countries in our sample set tariffs about 9 percentage points higher in goods with medium or high market power relative to those with low market power; 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' tariffs prior to their WTO membership so they are set in a unilateral, noncooperative way. However, we also analyze the role of market power in shaping a subset of trade policies that are determined noncooperatively by the United States, a large member of the WTO. The United States sets nontariff barriers and statutory tariffs (i.e., rates it applies to some non-WTO members) with few or no restrictions from the WTO. We find that market power is also an important determinant of these trade policies that the United States sets unilaterally. Interestingly, we find no such effect on those US tariffs set according to WTO rules. This finding is broadly consistent with Bagwell and Staiger's theory of the WTO, and it suggests that market power would play an important role for *all* US trade policies if they were set noncooperatively, e.g., in the absence of the WTO. More generally, the results for the United States show that the importance of the terms-of-trade motive extends to 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 II, we describe the estimation methodology for the elasticities. In Section III, we describe the data and assess the validity of the elasticity estimates. We present the estimation results of the impact of market power on trade policy in Sections IV and V, and conclude in Section VI.

I. 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.

A. *Optimal Tariffs: The Benchmark Case*

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. 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 this utility, the demand for each good g is simply a function of its own price, i.e., $c_g = c_g(p_g)$. Social welfare is then the sum of the individual indirect utilities, which includes income and consumer surplus:[3]

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

To determine income, we employ the standard assumptions in the leading endogenous trade policy models, e.g., Gene Grossman and Elhanan Helpman (1994, 1995). First, the numeraire is

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

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 nonnumeraire goods are produced under constant returns to scale using labor and one factor specific to the good. This means that 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. If we normalize the population to be one and recall the wage is also unity, we can rewrite social welfare as

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

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

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

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, i.e., $p_g(\tau_g), p_g^*(\tau_g)$.[4]

A government choosing the tariff to maximize (3) will set it according to the following first-order condition:[5]

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

The first term represents the domestic distortion caused by the negative impact of tariffs on import levels. 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,[6]

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

B. *Optimal Tariffs: Extensions*

The positive relationship between protection and the inverse elasticity, ω_g, extends to more general settings. Here we highlight a few points. The separability assumption in our model implies that the tariffs in (6) do not reflect any monopoly power in the export sector. The bottom line from studies that consider market power in the export sector is that market power may create an additional motive for the use of import tariffs (c.f. Jan de V. Graf 1949–1950 and Fernando E. Alvarez and Robert E. Lucas 2007; Daniel Gros (1987), shows this is the case even for "small" countries when products are differentiated), but this additional motive does not

[4] 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 that is independent of prices.

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

[6] 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.

eliminate the first-order incentive to impose higher tariffs in sectors in which imports are supplied less elastically.

The positive relationship between tariffs and inverse elasticities also holds *even if* the government's objective is not social welfare maximization. For example, Grossman and Helpman (1995) extend their political contributions trade model to the large country case. The noncooperative 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 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 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 the government places *no* weight on social welfare, which we can immediately see by setting a equal to zero and noting that the second term in (7) remains finite. In sum, even though the terms-of-trade motive for tariffs is often associated with a welfare-maximizing government, such a relationship can also arise even if governments care *only* about lobbies' contributions.[7]

In equation (7) 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.[8] 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 of our working paper (available on request), 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 happens 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.[9]

[7] In this setting, this occurs because lobbies' contributions account for all the costs and benefits of the tariffs they bid on, including the terms-of-trade gain the lobbies reap via the redistributed tariff revenue.

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

[9] As we prove in that 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.

II. 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. In fact, most estimates 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).[10]

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:

$$\Delta^{k_{ig}} \ln s_{igvt} = -(\sigma_{ig} - 1)\,\Delta^{k_{ig}} \ln p_{igvt} + \varepsilon_{igvt}^{k_{ig}}; \tag{8}$$

$$\Delta^{k_{ig}} \ln p_{igvt} = \frac{\omega_{ig}}{1 + \omega_{ig}}\,\Delta^{k_{ig}} \ln s_{igvt} + \delta_{igvt}^{k_{ig}}. \tag{9}$$

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}. Thus, 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 over time. 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. One important shock to supply is bilateral exchange rate *changes* between countries i and v. We can see how exchange rates would enter into δ_{igvt} by rewriting the domestic price as

$$p_{igvt} = (1 + \tau_{igvt})\, e_{ivt}\, p_{igvt}^{*},$$

where τ_{igvt} is some ad valorem trading cost (e.g., tariffs), and e_{ivt} is the bilateral exchange rate between i and v. In this case, k-differencing would produce

$$\Delta^{k_{ig}} \ln p_{igvt} = \Delta^{k_{ig}} \ln(1 + \tau_{igvt}) + \Delta^{k_{ig}} \ln e_{ivt} + \Delta^{k_{ig}} \ln p_{igvt}^{*},$$

so the export supply error, $\delta_{igvt}^{k_{ig}}$, contains the bilateral exchange rate shock. Since these are frequent and large, they are likely to be a more important source of variation than shocks to relative trade costs—a point that we discuss further below.

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 assumed to be uncorrelated, i.e.,

[10] 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 United States. Irwin (1988) and John Romalis (2007) report both elasticities. However, because they are estimated at the aggregate level and for only two countries (the United Kingdom and United States, respectively), they cannot be used to estimate the impact of market power on tariffs.

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

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

$$\text{where } \theta_{ig1} = \frac{\omega_{ig}}{(1+\omega_{ig})(\sigma_{ig}-1)}, \theta_{ig2} = \frac{\omega_{ig}(\sigma_{ig}-2)-1}{(1+\omega_{ig})(\sigma_{ig}-1)}, \text{ and } u_{igvt} = \frac{\varepsilon_{igvt}^{k_{ig}} \delta_{igvt}^{k_{ig}}}{\sigma_{ig}-1}.$$

Note that the new error term, u_{igvt}, is correlated with the "independent" variables in equation (10) that depend on prices and expenditure shares. However, Feenstra (1994) shows that a consistent estimator of $\boldsymbol{\theta}_{\mathbf{ig}} = (\theta_{ig1}, \theta_{ig2})$ 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 $Y_{igvt} = (\Delta^{k_{ig}} \ln p_{igvt})^2$, $X_{1,igvt} = (\Delta^{k_{ig}} \ln s_{igvt})^2$, $X_{2,igvt} = (\Delta^{k_{ig}} \ln p_{igvt} \Delta^{k_{ig}} \ln s_{igvt})$, and the bars on top of these variables denote their time averages (the t subscript is dropped). The independence of errors assumption implies that $E_v(\bar{u}_{igv}) = 0$. 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, the true underlying elasticities are exactly identified. 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.

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 this fails, 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.[11]

The precision for the typical elasticity is obtained by bootstrapping. We resampled 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 is similar to the one used in the estimation of the actual elasticities.

Several features of this estimation strategy help us to avoid concerns that usually make it difficult to obtain consistent estimates. First, one might be worried that if countries impose tariffs in response to demand shocks, this might cause a correlation between demand and supply shocks.

[11] We present additional details about the specific computational procedure in our working paper.

However, if these tariff changes are implemented identically across exporters of any given good g (as they often are), then that effect is purged from the export supply error in our estimation since $\Delta^{k_{ig}} \ln(1 + \tau_{igvt}) = \Delta \ln(1 + \tau_{igvt}) - \Delta \ln(1 + \tau_{igk_{ig}t}) = 0$. As a result, such tariff changes will not affect the estimates. This also 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.[12]

Note that the double differencing is also useful in controlling for other factors that could otherwise induce a correlation of the error terms. For example, if some countries produce higher quality goods at higher cost, the time-differencing of the data will eliminate any correlation in the levels. Similarly, if the quality and cost of a good are rising, the time and k-differencing will eliminate any correlation between demand and supply as long as the trend in quality is common across exporters.

In the robustness section, we analyze if our results are sensitive to some of the 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, six to nine 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 of our working paper, we test and find evidence that supports this assumption.

III. Data, Descriptive Statistics, and Assessment of Elasticity Estimates

A. *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 noncooperative way. Since a major function of the General Agreement on Tariffs and Trade (GATT)/WTO is to allow countries to reciprocally lower their tariffs, possibly in order to internalize the terms-of-trade effects, we focus the test on policies that countries set prior to the constraints of GATT/WTO membership. In Section V, we provide additional evidence for a set of policies set noncooperatively by a WTO member.

[12] In an extreme version of the optimal tariff argument, we may expect countries to discriminate across all different exporters of the same good. This would entail very high administrative costs and thus is not the norm. The closest countries come to such discrimination is through preferential agreements. These agreements are not important for most countries in our sample. Some of them did, however, implement such agreements during the period for which we estimate elasticities. Those differential tariff changes are reflected only on the export supply error, $\delta^{k_{ig}}_{igvt}$, since the demand equation controls for the domestic price. Thus, such shocks do not invalidate our elasticity identification assumption unless there is some other simultaneous shock to relative demand. We address the possibility that such preferential tariff changes affect both the elasticity estimate *and* the nonpreferential tariffs by using instrumental variables in the tariff estimation section.

Table 1—Data Sources and Years

	GATT/WTO	Production data		Tariff data[a]	Trade data[b]
	Accession date	Source	Years		
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

[a] All tariff data are 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 are from COMTRADE. For Taiwan, data are 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 1990–1993.

[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 IVC, the pooled tariff results are robust to dropping the Czech Republic.

Our tariff data come from the TRAINS database, which provides data at the six-digit HS level. We focus on the 15 countries that report tariffs in at least one-third of all six-digit goods.[13] The set of countries and the years we use are reported in Table 1. Our sample includes a nonnegligible 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 1995 world average of $8,900. The 15 countries comprise 25 percent of the world's population and close to 20 percent of its GDP (in PPP terms). 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 nonnegligible countries such as Taiwan, Ukraine, Algeria, and Saudi Arabia.

The trade data are obtained from the United Nations Commodity Trade Statistics Database (COMTRADE). This database provides quantity and value data at six-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 cover the period 1994–2003. For Taiwan we use the United Nations Conference on Trade and Development (UNCTAD) TRAINS database since COMTRADE does not report data for this country.

B. *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

[13] Unfortunately, some non-WTO countries report this tariff data for only a small share of goods, making it impossible to make meaningful comparisons across goods. Our criteria were binding only for the Bahamas, Brunei, Seychelles, and Sudan.

TABLE 2—TRADE AND TARIFF DATA SUMMARY STATISTICS

	Trade data			Tariff data				
				Rate per four-digit HS				
	Number of varieties[a]	Number of HS4 goods	Median # of var. per HS4	Observations[b]	Mean	Standard Deviation	Median	Fraction of HS6 variation between HS4
Algeria	26,466	1,100	13	739	23.8	17.4	15.6	0.95
Belarus	24,440	1,172	12	703	12.4	7.8	10.0	0.94
Bolivia	18,592	1,064	9	647	9.8	0.8	10.0	0.63
China	63,764	1,217	33	1,125	37.9	26.0	30.3	0.93
Czech Republic	61,781	1,219	30	1,075	9.5	17.6	5.1	0.87
Ecuador	22,979	1,101	11	753	9.8	5.5	10.6	0.91
Latvia	33,790	1,128	17	872	7.3	10.5	1.0	0.90
Lebanon	34,187	1,109	15	782	17.1	14.8	15.0	0.87
Lithuania	34,825	1,159	17	811	3.6	7.4	0.0	0.90
Oman	20,482	1,107	10	629	5.7	8.7	5.0	0.76
Paraguay	15,430	1,049	7	511	16.1	11.3	14.0	0.91
Russian	66,731	1,187	34	1,029	10.7	11.0	5.7	0.95
Saudi Arabia	62,525	1,202	32	1,036	12.1	2.6	12.0	0.93
Taiwan	38,397	1,215	19	891	9.7	8.5	7.5	0.90
Ukraine	37,693	1,128	18	730	7.4	7.6	5.0	0.95
Median	34,187	1,128	17	782	9.8	8.7	10.0	0.91

[a] Varieties are defined as six-digit HS, exporting country pairs.
[b] Number of observations for which elasticities and tariffs are available.

become imprecise. Therefore, in estimating (8) and (9) we define a good, g, as a four-digit HS category and a variety, v, as a six-digit good from a particular exporter. Table 2 shows that the typical country has 1,100 four-digit categories with positive imports between 1994 and 2003. The typical good in the sample contains 17 HS6-country pairs. 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 (four-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 tariff value. Finally, since we estimate the elasticities at the HS4 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 HS4 variation has little impact since over 90 percent of the variation in tariffs for the typical country occurs across HS4 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 suggest 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 If we drop China, however, 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

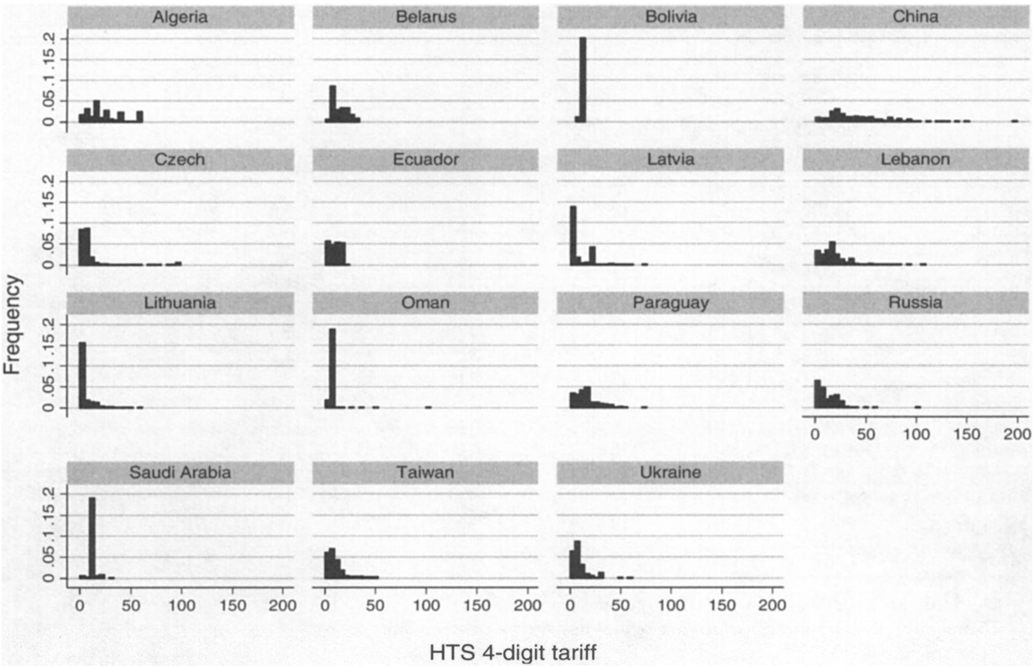

FIGURE 1. TARIFF DISTRIBUTION BY COUNTRY

within-country frequency distribution of tariffs at the four-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, such as Oman and Saudi Arabia. Moreover, we observe truncation and some bunching at the lower end of distribution, where about 9 percent 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 seem 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 policymakers are uncertain of 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 as tariffs approach 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 confirm or reject the optimal tariff theory expressed in a particular functional form, but rather to estimate the impact of market power on tariffs.

C. *Elasticity Estimates*

Since we conduct the analysis at the four-digit level for each country, we estimate over 12,000 foreign export supply elasticities—far too many to present individually. Therefore, in Table 3A

TABLE 3A—INVERSE EXPORT SUPPLY ELASTICITY STATISTICS

Statistic	Observations[a]	Median[b]			Mean		Standard 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	2.0	91	102	23	283	49
China	1,125	0.4	2.1	80	92	17	267	35
Czech Republic	1,075	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	3,536	21
Paraguay	511	0.4	3.0	153	132	67	315	169
Russia	1,029	0.5	1.8	33	48	8	198	18
Saudi Arabia	1,036	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

[a] 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 HS4 goods for which elasticities were computed.

[b] The median over the "low" sample corresponds to the median over the bottom tercile of inverse elasticities. Medium and high correspond to the second and third terciles.

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, i.e., 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 thirty-third 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. The problem of outliers can be seen from the fact that when we trim the top decile of the sample in Table 3A, the means fall by almost an order of magnitude, down to 13. The same is true for the standard deviation. Since the standard errors are nonspherical, we assess the precision of the estimates via bootstrapping. In Table 3B, we report results from resampling the data and computing new estimates for each of the elasticities 250 times.[14] Table 3B indicates that the imprecision of the estimates appears to be most severe 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. 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.

[14] 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.

TABLE 3B— BOOTSTRAPPED STATISTICS FOR INVERSE EXPORT SUPPLY ELASTICITIES

Statistic	Low		Medium or high	
Sample	Median	Confidence interval[a]	Median	Confidence interval[a]
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 Republic	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 , 60]
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]

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

[a] 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 $1 - 2\alpha$ confidence interval are obtained via the bias-corrected percentile method (Bradley Efron 1981) using $\alpha = 0.1$.

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, we 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 $[0.8, 3]$ for the median *medium* market power good and 33 $[3, 53]$ 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.[15] We will be conservative and try to distinguish only between low versus medium or high market power goods. The data clearly allow us to 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 toward 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 address measurement error explicitly.

[15] 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.

D. *Assessment of Elasticity Estimates*

We now turn to the question of whether our estimates themselves are plausible. In our estimation framework, we rely only on the relative levels of market power across goods, so the focus of our tests will be on the relative magnitudes, although we will discuss the levels at the end of this section. We explore the "reasonableness" of the variation in elasticities in three ways. First, we check whether the 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 is often stressed by the trade literature, or if they are in more remote regions.[16]

The motivation for the tests is clearer if we note that the residual supply of exports faced by importer i, m^*_{igv}, is 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 importer i, $1/\omega_{igv}$, should be an increasing function of both the exporter's supply elasticity, denoted by λ^*_{gv}, a weighted average of other countries' 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

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

We 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, ω_{igv} and ω_{jgv} should be similar because countries other than i and j also consume the good and these demand elasticities will enter identically in the export supply equation. 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 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. 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.[17]

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., US soybeans, and as a consequence the price falls, then other countries will substitute toward that good and away from other sources of supply (e.g., demand less of other countries' soybeans), so the equilibrium price decline will be minimal. Such substitution

[16] In the working paper, we also show that the elasticities are consistent with the implied pass-through coefficients in other studies. These results are available upon request.

[17] We use a log specification to minimize the influence of the outliers and because the estimation procedure for the elasticities cannot yield nonpositive estimates.

Table 4—Correlation of Inverse Export Supply Elasticities across Countries

	Log inverse export supply			
Dependent variable: Statistic	Beta	Standard error	R^2	Number of observations
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	1,125
Czech Republic	0.61	(0.05)	0.12	1,075
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	1,029
Saudi Arabia	0.48	(0.06)	0.08	1,036
Taiwan	0.31	(0.08)	0.02	891
Ukraine	0.83	(0.07)	0.17	730
Median	0.70	(0.07)	0.12	782

Note: 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 8,734, less than the full sample since not all HS4 can be uniquely matched to Rauch's classification. The number for the mean regression is 7,927 because we trim the top decile. The pattern of results with the top decile is similar but with higher values.

is much less likely for specialized or differentiated goods such as locomotives, aircraft, or integrated circuits because no two exporters produce the exact same good. Thus, we conjecture that countries have more market power in differentiated goods than commodities.

James E. 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 three times larger than reference goods and five 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

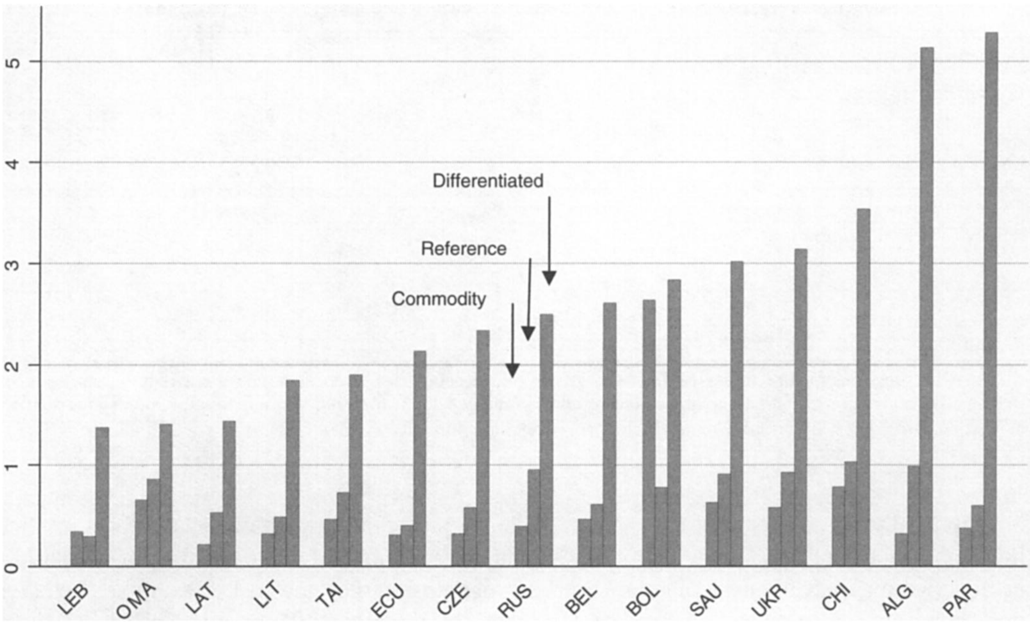

FIGURE 2. MEDIAN INVERSE ELASTICITIES BY PRODUCT TYPE
(*Goods classified by Rauch into commodities, reference priced products, and differentiated products*)

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 locomotives and integrated circuits is more than double the sample median. These are all differentiated goods for which it is more likely that even 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 subsample 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 four-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.[18] 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.[19]

[18] This is consistent with the results in James R. Markusen and Randall M. 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 the United States than for Canada.

[19] 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. The within *R*-square for GDP is also small, which explains why tariffs and GDP in our sample do not have a

TABLE 6—INVERSE EXPORT SUPPLY ELASTICITIES, GDP, REMOTENESS, AND IMPORT SHARES

Dependent variable	Log inverse export supply		
Log GDP	0.17 (0.04)	0.18 (0.03)	
Log remoteness		0.40 (0.15)	
Share of world HS4 imports			7.19 (1.48)
Observations	12,343	12,343	12,343
R^2	0.26	0.26	0.25
R^2 within	0.01	0.02	0.00

Notes: All regressions include four-digit HS fixed effects (1,201 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 $1/(\Sigma_j \text{GDP}_j/\text{distance}_{ij})$. The share of world imports is calculated in 2000.

Empirically, we know that trade volume falls off quite rapidly with distance.[20] This implies that some goods are traded only 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 share of demand for certain regionally traded goods, such as Chilean cement, and it is this elasticity that we estimate. 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.

Finally, consider 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. First, as noted above, there are still large trade costs segmenting markets. Second, although we do find that countries have almost no market power in homogenous goods (those that Rauch (1999) defines as commodities), those goods make up only about 10 percent of the tariff lines in the sample. About 60 percent of the HS4 goods in the sample are differentiated with the remaining 30 percent classified as reference priced.

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.

robust positive correlation (e.g., it disappears once we drop China), but tariffs and inverse elasticities do, as we show in the next section.

[20] According to James E. Anderson and Eric van Wincoop's review of the literature, "the tax equivalent of 'representative' trade costs for industrialized countries is 170 percent" (2004, 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 nonbordering country if one considers the decay due to distance alone (c.f. Limão and Anthony J. Venables 2001).

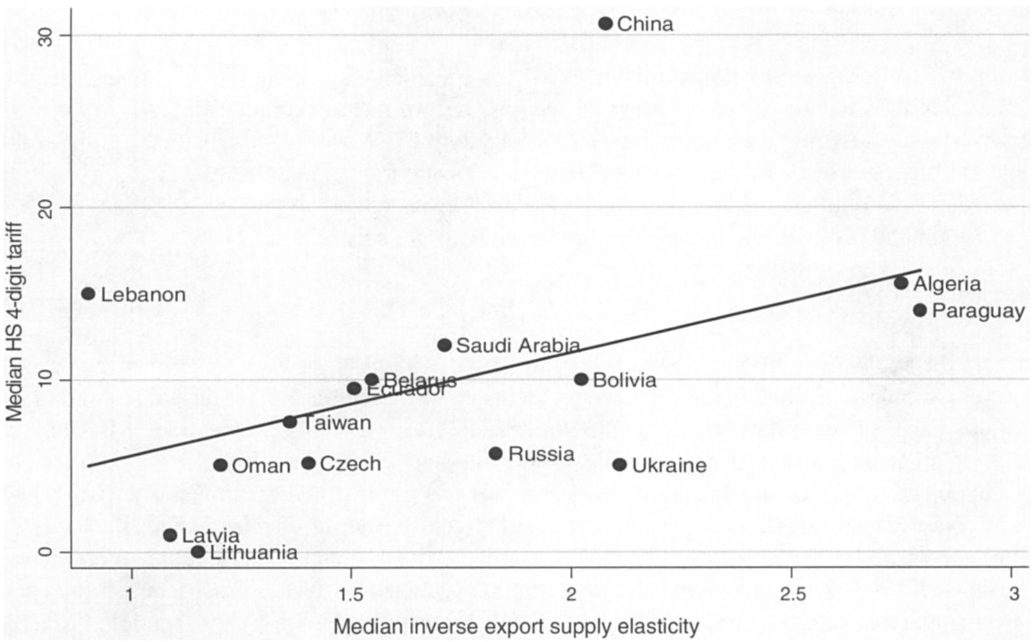

FIGURE 3. MEDIAN TARIFFS AND MARKET POWER ACROSS COUNTRIES

IV. Estimating the Impact of Market Power on Tariffs

A. *Cross-Country Evidence*

Before turning to the regression evidence, we will examine a crude cross-country version of the theory. 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 particular income level. The positive relationship between median tariffs and median elasticities is also statistically significant.[21] Of course, there are many reasons to be wary of this relationship, starting with the small number of observations. Fortunately, the vast quantity of country-good data underlying this plot can be used to examine the relationship more carefully, and we do so in our original working paper where we confirm its robustness.

The result we have presented thus far is suggestive but still far from convincing. Expressing the tariff purely in terms of an aggregate 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.

[21] 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).

B. *Baseline Results*

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 subsequent subsections we provide a variety of robustness checks and further augment the model to include two specific prominent motives for protection: tariff revenue and lobbying.

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 four-digit good, g, as does the market power variable, ω. 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 $f(\cdot)$. Since our main objective is to establish whether market power is a significant determinant of 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. Many of the latter factors, however, 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.[22] 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. Any given industry can, however, 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 focus on primarily 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 trade policy models focus on industry-level determinants for which we can control. In any given country, however, 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 we control for all relevant ones. Thus, our main strategy for addressing omitted product variables in this section is to use instrumental variables. Subsequently, 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. Since the results in columns 1–3 are qualitatively similar to the ones in columns 4–6, we discuss the latter.[23] The

[22] The industry that contains the good is defined to be one of the 21 sections of the Harmonized Tariff Schedule, e.g., textiles, chemicals, plastics, etc.

[23] The comparable three specifications with industry-by-country effects are also similar and available upon request.

TABLE 7— TARIFFS AND MARKET POWER ACROSS GOODS (WITHIN COUNTRIES): OLS AND TOBIT ESTIMATES

Dependent variable	Average tariff at four-digit HS (%)								
Fixed effects	Country			Country and industry					
Estimation method	OLS	OLS	OLS	OLS	OLS	OLS	Tobit	OLS[a]	OLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Inverse exp. elast.	0.0003			0.0004					
	(0.0001)			(0.0004)					
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 Republic	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	12,333	12,333	12,333	12,333	12,333	12,333	12,333	12,333	12,333
Number of parameters	16	16	16	36	35	36	35	38	36
Adj. R^2	0.61	0.61	0.61	0.66	0.66	0.66			0.66

Notes: Standard errors in parentheses (all heteroskedasticity robust except Tobit). Industry dummies defined by section according to Harmonized Standard tariff schedule.

[a] Optimal threshold regression based on minimum RSS found using a grid search over 50 points of the distribution of inverse exp. elast. (from first to ninety-ninth percentile in intervals of two). Optimal threshold is fifty-third percentile. Accordingly, med hi equals one above the fifty-third percentile and zero otherwise. Bruce E. 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.

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 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 is 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.[24]

A parsimonious way to address the skewness of market power and its nonlinear 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 one if it is above the thirty-third 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 sixty-sixth percentile of inverse export elasticity) or medium (between thirty-third and sixty-sixth). 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 now turn to estimates that employ 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. When we employ the continuous measures, however, 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 error. When assessing the elasticities, we showed that our estimates clearly distinguish between goods where a country has low versus medium or high market power. 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 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 last 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

[24] It estimates a slope of 1.9 when market power is below the estimated threshold (fifty-third percentile), which is considerably larger than the slope above it. The threshold in a similar specification without industry effects is at the thirty-third percentile.

TABLE 8—TARIFFS AND MARKET POWER ACROSS GOODS (WITHIN COUNTRIES): IV ESTIMATES

Dependent variable	Average tariff at four-digit HS (%)								
Fixed effects	Country			Country and industry			Industry by country		
Estimation method	IV GMM (1)	IV GMM (2)	IV GMM (3)	IV GMM (4)	IV GMM (5)	IV GMM (6)	IV GMM (7)	IV GMM (8)	IV GMM (9)
Inverse exp. elast.	0.040 (0.027)			0.089 (0.055)			0.075 (0.028)		
Mid and high inv. exp. elast.		3.96 (0.76)			8.88 (1.18)			9.07 (1.08)	
Log(1/export elasticity)			0.75 (0.15)			1.71 (0.23)			1.73 (0.21)
Observations	12,258	12,258	12,258	12,258	12,258	12,258	12,258	12,258	12,258
No. of parameters	16	16	16	35	35	35	284	282	283
1st stage F	5	1649	1335	2	653	517	3	691	544

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

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 percent level. 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 IVF.[25]

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. Given its prominence in the basic theoretical prediction, however, 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.[26]

C. *Individual Country Results*

To carefully establish the tariff determinants of any given country requires its own paper. We want to determine, however, 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. Yet we cannot ignore obvious issues such as the bunching of tariffs in Bolivia, Oman, and Saudi Arabia.

[25] 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 ninetieth percentile in each country. However, when we reestimate the IV without those observations, we obtain very similar estimates for β.

[26] 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. If we drop these countries, the estimates become more precise and increase in magnitude in the dummy and semi-log specifications.

For the other 12 countries, we still employ the IV approach with industry effects and estimate the unrestricted version of (13) for each country.[27]

Table 9 presents the IV results by country using the semi-log specification. The first two columns reproduce the pooled results from Table 8 for comparison. The estimate is positive for each and every one of the 15 countries. It is also statistically significant at the 5 percent or 1 percent level for all but two.[28] 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 subsample of 12 countries.[29]

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, which is close to the mean over the country estimates, 0.17, as well as the value obtained for the typical country, 0.16. The range of elasticity estimates across countries is fairly narrow 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.

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.[30] As we pointed out in the previous paragraph, however, 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 is that in this sample average tariffs are higher for larger countries, as the theory would predict.

D. *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 regressions, 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 HS4 addresses this concern, and we verified that it does not change the significance of the results.[31]

[27] Tariffs in those three countries 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.

[28] One of them is Saudi Arabia, where we did not expect a precise estimate anyway. The other is the Czech Republic which, 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 (available upon request).

[29] Table 9, panel B, of Broda, Limão, and Weinstein (2006) presents the analog using the categorical variable. They are qualitatively similar to those presented in the semi-log specification in terms of sign and significance. 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.

[30] That is, in the general specification we model $\beta_i = \beta + \beta^L \times 1(\text{Size}_i)$, where the indicator variable 1(Size) is one if country i is above the sixty-sixth percentile in terms of size. We defined size as either GDP in 1996 or GDP adjusted by "regional market size," i.e., divided by an inverse distance weighted average of other countries' GDPs. Both measures identified the same six countries as the largest. 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).

[31] We also analyzed an alternative instrument with more variation over countries. The theory and our initial regressions of the determinants of market power suggest using a measure of country size to help predict market power in the first stage. Since the regression already includes country effects, we interact a measure of country size with our original

TABLE 9A—TARIFFS AND MARKET POWER ACROSS GOODS BY COUNTRY: IV ESTIMATES

Dependent variable	Average tariff at four-digit HS (%)									
Fixed effects	Ind. by Country	Ind. by Country	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
Sample	All	Exc. Bol., Om., Sau.	China	Russia	Taiwan	Saudi Arabia	Ukraine	Czech.	Algeria	Belarus
Log(1/export elasticity)	1.73	2.11	7.60	2.42	1.98	1.75	0.71	0.16	5.40	2.28
	(0.21)	(0.25)	(1.78)	(0.61)	(0.78)	(2.55)	(0.26)	(0.24)	(0.97)	(0.52)
Observations	12,258	9,952	1,089	1,021	841	1,031	685	1,000	739	703
No. of parameters	283	227	20	18	20		18	20	18	18
First stage F	544	477	45	44	12	7.6[a]	48	61	60	40
Mean tariff (%)	13.4	14.2	38.2	10.3	8.9	12.2	5.8	5.5	23.8	12.4
Elasticity (at mean)	0.13	0.15	0.20	0.23	0.22	0.14	0.12	0.03	0.23	0.18

TABLE 9B— TARIFFS AND MARKET POWER ACROSS GOODS BY COUNTRY: IV ESTIMATES

Dependent variable	Average tariff at four-digit HS (%)								
Fixed effects	Ind.		Ind.	Ind.	Ind.		Ind.		
Estimation method	IV GMM	IV Tobit	IV GMM	IV GMM	IV GMM	IV Tobit	IV GMM		
Sample	Ecuador	Oman	Paraguay	Lithuania	Lebanon	Bolivia	Latvia	Mean	Median
Log(1/export elasticity)	1.55	0.60	2.44	0.83	2.41	0.79	1.41	2.15	1.75
	(0.34)	(0.18)	(0.67)	(0.27)	(0.54)	(0.36)	(0.60)		
Observations	753	628	510	768	754	647	868		
No. of parameters	19		17	19	20		19		
First stage F	45	3.7[a]	33	35	52	8.96[a]	37		
Mean tariff (%)	9.8	5.7	16.0	2.3	15.0	9.8	7.0		
Elasticity (at mean)	0.16	0.11	0.15	0.36	0.16	0.08	0.20	0.17	0.16

Notes: Standard errors in parentheses (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 five countries, and even then only slightly since the criteria drop only 3–7 percent of their observations (those with tariff values more than three times the interquartile range above the seventy-fifth percentile or below the twenty-fifth). Bolivia, Oman, and Saudi Arabia are estimated via an IV Tobit procedure to account for the fact that a large fraction of their observations is censored from above and/or below. Given the lack of variation in their tariff within industries, their estimation does not include industry dummies.

[a] z-stat of the instrument in the first stage of IV Tobit.

We can relax the assumption of common elasticities across sets of exporters of a given good to country i. This is costly, however, 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 do 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

instrument. The magnitude and significance of the results for the three baseline specifications in columns 7, 8, and 9 of Table 8 are quite similar when we employ this alternative instrument that varies over both countries and goods.

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 rerun the baseline results in columns 7–9 of Table 8 with each set of estimates. Both sets yield a positive and significant 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.

E. *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. As noted before, much of the theory and empirical evidence on trade protection focuses on industry-level determinants, so it is difficult to know for which *product* characteristics we should control. Thus, we focus on two prominent motives for protection that may 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. It is simple to show, however, that if governments use tariffs to raise revenues they would impose higher tariffs on goods with lower import demand elasticities, σ_{ig}. This occurs because governments obtain more revenue and impose lower consumption distortions when they impose a given tariff on a good with a lower import elasticity. 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 zero for product g if its value of $1/\sigma_{ig}$ is in the bottom tercile in country i, and one 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. This indicates that these countries apply tariffs that are 9 percentage points higher in sectors in which they have market power. The estimate for the semi-log is also statistically identical to the baseline. 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.

The specific political economy factors that are relevant for the tariff structure can also differ across these countries. So, we include a political economy variable that is central to 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

TABLE 10— MARKET POWER VERSUS TARIFF REVENUE OR LOBBYING AS A SOURCE OF PROTECTION

Dependent variable	Average tariff at four-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	12,258	12,258	12,258	12,258	5,178	5,178
No. of parameters	282	283	283	284	132	133
First stage F (market power)	691	544	370	312	171	129
First stage F (other)	na	na	102	144	131	188

Notes: Standard errors in parentheses (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. These data are not available for mining and agricultural products.

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. This is available for all these countries only at the ISIC three-digit data from the United Nations Industrial Development Organization (UNIDO) industrial database. So z_{ig} can be interpreted as country i's average penetration for the goods in that ISIC three-digit category. Since we divide this by the import demand elasticity, which varies by HS4, the lobbying variable also varies at the HS four-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 zero for the lower tercile of z_g/σ_g in that country, and one 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 statistically indistinguishable from the baseline results. The same conclusion holds if we consider the semi-log specification. Note also that the reason 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 only by itself, but also relative to this important alternative explanation.

Table 11—Market Power and Lobbying: IV Estimates by Country

Panel A: Semilog										
Dependent variable	Average tariff at four-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	5,178	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	39	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 four-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 high 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	5,178	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-stage F: log(imp. pen/imp. el)	131	37	19	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 parentheses (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 fewer than 36 observations in any of these countries (those with tariff values more than three times the interquartile range above the seventy-fifth percentile or below the twenty-fifth). 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. The first-stage statistic in this case refers to the z-statistic of the relevant instrument.

Table 11 shows that these results are not driven by a single country. In all countries for which we have production data, market power has a positive effect on tariffs, which is statistically and economically significant.

F. *Quantification*

We now quantify the impact of market power in explaining tariffs and the implication of removing this motive for tariff setting on prices received by exporters.

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 8)	Subsample[a] (Table 8)	Typical country[b] (Table 9)	Pooled (Table 10)	Typical country (Table 11)
	β	1.7 pp	2.1 pp	1.8 pp	1.9 pp	1.5 pp
Log(1/export elasticity)	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.($\ln\omega$)/γ×s.d.($\ln z/\sigma$))	.	.	.	1.6	1.5
	β	9 pp	11 pp	9 pp	10 pp	8 pp
Mid and high inv. exp. elast.	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.

[a] Equivalent of specifications in columns 8 and 9 of Table 8 excluding Bolivia, Oman, and Saudi Arabia.

[b] Estimates for bottom rows are obtained from the equivalent of specifications in Table 9 using the categorical variable.

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. The pooled regression in Table 10 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 one-standard-deviation change on 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: 92 percent of the average in the baseline case and 97 percent in the lobbying augmented model. Given that the effect applies to two-thirds of the tariff lines, the estimates imply that if these countries did not exert their market power, the implied tariff reductions would be around 60 percent of existing levels. This is much larger than the Uruguay Round's 25 percent tariff reduction target for developing countries.

If countries possess and exert market power, then their tariffs should depress the price received by foreign exporters. Thus, reducing tariffs increases those prices. Our estimates allow us to calculate the exporter's price increase that results when an importer, i, reduces its tariffs. 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 1930s 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 its market power, or reflect it only partially. The counterfactual also provides an estimate of the impact of a country leaving the agreement and reexerting its market power, which is obtained 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 $p^*_{igv} = p^*_{igv}(\tau_{ig}(\omega_{ig}), .)$, i.e., a function of the tariff it faces and other parameters that are omitted. The percent increase in that price if an importer were to start treating a given good with medium or high market power as if it had low market power would then be

$$d\ln p^*_{igv} = \frac{d\ln p^*_{igv}}{d\ln(1+\tau_{ig})\big|_{\omega_{ig}^{med\_hi}}} \times [\ln(1+\tau_{ig})\big|_{\omega_{ig}^{low}} - \ln(1+\tau_{ig})\big|_{\omega_{ig}^{med\_hi}}] \times 100, \tag{14}$$

$$\approx -(\zeta_{ig}-1)^*\beta_i,$$

where ζ_{ig} is simply the domestic pass-through, i.e., the effect on the domestic price of good g in country i of a 1 percent increase in country i's tariff factor. Since $p_{igv} = (1+\tau_{ig})p^*_{igv}$, we have $\zeta_{ig} = 1 + (d\ln p^*_{igv}/d\ln(1+\tau_{ig}))$, so $\zeta_{ig} - 1$ 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.[32]

In the appendix of our working paper, we show that $\zeta_{ig} = 1/(1+\omega_{ig})$ in our framework. The typical good in the middle tercile in terms of market power has an inverse elasticity of 1.6, implying a domestic pass-through of 0.4, which is similar to the values in that literature. Therefore, the impact on exporter prices, $\zeta_{ig} - 1$, equals $-0.6\,(= 0.4 - 1)$. Multiplying this by the estimated tariff reduction due to treating those goods as low market power goods (9 percentage points in Table 8, column 8), we obtain a 5 percent increase in the price received by the exporters. By construction, this applies to a third 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 in our working paper, the effect for Russia and Taiwan is 6–10 percent and for China it is particularly large, 17–25 percent. 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. Conversely, if these countries

[32] The approximation arises because in equation (13) we use tariffs rather than ln $(1+\tau_{ig})$, 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 $\zeta_{ig} - 1$ 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 an additional import demand for v due to substitution away from exporters of other varieties of g not in the WTO.

were to abandon an agreement and reexert 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.

V. Market Power and Trade Barriers in a Large Developed WTO Member

Our focus on non-WTO members is motivated by the theory that predicts that a country's tariffs will reflect its market power when it acts noncooperatively. 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 most favored nation (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; 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 US protection and use nontariff barriers (NTBs) as the measure of its noncooperative trade policy (e.g., Goldberg and Maggi 1999; Kishore Gawande and Ursee Bandyopadhyay 2000). Several of these NTBs—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 six-digit goods within each four-digit classification that contain an NTB. We complement this by using the ad valorem equivalent recently estimated by Hiau L. Kee, Alessandro Nicita, and Marcelo Olarreaga (forthcoming).

Second, we examine "statutory rates"—the tariffs the United States applies to countries to which it does not grant MFN status. Statutory tariffs are set noncooperatively, which is apparent from their high levels and the targeted countries.[33] Successive rounds of trade negotiations opened a large gap between these rates and the MFN rates the United States negotiates with and applies to WTO members. The average US 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

[33] 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.

TABLE 13— MARKET POWER AND LOBBYING AS A SOURCE OF PROTECTION IN THE US

Panel A: Nontariff barriers

Theory	*Market power*				*Market power and lobbying*			
Fixed effects	Industry				Industry			
Estimation method	IV Tobit				IV Tobit[b]			
Dependent variable	Coverage ratio (HS4)[a]		Advalorem equiv. (HS4, %)		Coverage ratio (HS4)		Advalorem equiv. (HS4, %)	
	(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[c]	804	804	804	804	708	708	708	708
Number of parameters	17	17	17	17	17	17	17	17
First stage z-stat (market power)	7.1	6.6	7.1	6.6	6.2	5.3	6.2	5.3
First 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[b]			
Dependent variable	Non-WTO (HS4, %)		WTO (HS4, %)		Non-WTO (HS4, %)		WTO (HS4, %)	
	(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[c]	870	870	869	869	775	775	774	774
Number of parameters	20	20	20	20	21	21	21	21
First stage z-stat (market power)	7.3	7.1	7.3	7.1	6.0	5.3	6.0	5.3
First 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 parentheses. Industry dummies defined by section according to the Harmonized Standard tariff schedule.

[a] Coverage ratio is defined as the fraction of HS6 tariff lines in a given HS4 category that had an NTB. Since it varies between zero and one 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.

[b] 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.

[c] 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 nontariff 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.

insight into how the United States sets tariffs noncooperatively. Thus they are a useful complement to US NTBs, which apply to many countries.

Our estimation strategy is similar to the one used thus far. We estimate the relevant elasticities for the US using the same procedure, and use them to estimate equation (13), including industry dummies and instrumenting for the remaining variables. The structure of US production, trade, and demand differs in important ways from those of the developing countries we analyzed. Thus, we construct instruments for US elasticities and import penetration ratios as before, but use data from large developed countries: Canada, France, Germany, Japan, and the United Kingdom.[34]

In Table 13 we report the results for the United States. The results in panel A show that the United States sets significantly higher NTBs in products where it has more market power. This is true if we measure NTBs by the coverage ratio (columns 1, 2, 5, and 6) or the ad valorem equivalent (columns 3, 4, 7, and 8). The magnitudes for the ad valorem specification are large since products with NTBs have large tariff equivalents—about 18 percent for the typical HS4 with an NTB. These NTBs affect a large number of products—a quarter of the HS4 lines in the sample.[35] The results are robust to including the 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 United States sets noncooperatively. 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 US MFN rates (columns 3, 4, 7, and 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 US trade policy in areas not covered by the WTO. In other words, when the US can set trade barriers noncooperatively, it takes market power into account. This strongly suggests that market power would play an important role for *all* US trade policy if it were set noncooperatively, e.g., in the absence of the WTO.

VI. 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. This motive continues to play a key role in most theoretical models of trade policy, but there has been considerable disagreement about its practical importance as an empirical determinant of tariffs. Despite this, no one had thus far tested whether countries set higher tariffs in goods in which they have more market power.

Part of the reason for the absence of such tests was that many economists simply assumed that most countries are small, i.e., do not have market power in trade. One of the contributions of this work is to demonstrate that this assumption is not correct. This is likely to have important implications in areas like computable general equilibrium modeling where small country assumptions are often the rule. One additional area that deserves further research surrounds how the

[34] The US elasticity estimates are reasonable when we use the criteria previously applied. For example, they are strongly correlated with those of these five countries. Moreover, the typical inverse elasticity is highest for differentiated (1.6) than reference priced goods (0.55) or commodities (0.41).

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

elasticities we estimate change over time. Our elasticities are computed over a one-year horizon, and it is possible that one might obtain different elasticity values over longer time horizons. As long as the ranking across goods remained unchanged, however, our main tariff results would hold.

Our paper is the first to provide evidence that when countries are not subject to constraints such as those imposed in the WTO, they set 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 also show that the impact of market power on tariffs is economically significant. It is of the same magnitude as the average tariffs of the countries we examine, 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 and suggest potentially large gains from *reciprocal* tariff liberalization agreements such as the WTO.

We also find that market power strongly affects the noncooperative trade policies of a large developed country, the United States. 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 to WTO members, and understanding its implications for trade policy is essential. One such implication arises from our finding that market power significantly affects trade policies not subject to WTO constraints, but not those policies where such constraints are present (at least for the United States). This indicates that the WTO plays a quantitatively important role in reducing protection, unlike what is suggested by recent research.[36]

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