INSECURITY AND THE PATTERN OF TRADE: AN EMPIRICAL INVESTIGATION

James E. Anderson and Douglas Marcouiller*

Abstract—Corruption and imperfect contract enforcement dramatically reduce international trade. This paper estimates the reduction using a structural model of import demand in which insecurity acts as a hidden tax on trade. We find that inadequate institutions constrain trade as much as tariffs do. We also find that omission of indices of institutional quality biases the estimates of typical gravity models, obscuring a negative relationship between per capita income and the share of total expenditure devoted to traded goods. Finally, we argue that cross-country variation in the effectiveness of institutions and the consequent variation in the prices of traded goods offer a simple explanation for the stylized fact that high-income, capital-abundant countries trade disproportionately with each other.

I. Introduction

POPULAR accounts of globalization skirt a provocative question: Why is there so little international trade? Why is there less North-South trade than one would expect given differences in relative factor endowments? Why do international borders constrain exchange much more than can be explained by tariffs, quotas, and formal impediments to trade?

We contend that trade is reduced by hidden transactions costs associated with the insecurity of international exchange: contracts may not be enforceable across jurisdictional boundaries, bribes may be extorted by customs officials, and shipments may even be hijacked. Anecdotal evidence of insecurity is easy to find. Table 1, for example, shows the ranking in order of importance of "the obstacles for doing business" as reported in a 1996 World Bank survey of 3,685 firms in 69 countries (Brunetti, Kisunko, & Weder, 1997, p. 59). It is not surprising that firms complain about taxes; it is remarkable, however, that corruption ranks as the second most important obstacle to business worldwide, with crime and theft not far behind.

This paper develops a model of import demand in an insecure world. Our structural model is grounded in a multicountry generalization of the model of predation presented by Anderson and Marcouiller (1998); a similar structure can be derived from the model of incomplete contract enforcement developed by Anderson and Young (1999).

Fitting our structural model to the data, we find that trade expands dramatically when it is supported by strong institutions—specifically, by a legal system capable of enforcing commercial contracts and by transparent and impartial formulation and implementation of government economic policy. We estimate, for example, that, if the indices of institutional quality associated with the Latin American countries in our sample were to rise to the levels associated with the European Union, Latin American trade would expand by 30%, other things equal. This expansion is roughly equivalent to what we would expect from the reduction of Latin American tariffs to U.S. levels.

Empirical work that ignores the security of exchange suffers from an important omitted-variables bias. Our model shows that the share of expenditure devoted to traded goods falls as income per capita rises. When institutional variables are excluded from the regression, however, as in most of the existing gravity literature, traded goods expenditure shares appear to be positively associated with income per capita. The bias is due to the positive correlation of income per capita and institutional quality.

The stylized fact that high-income, capital-abundant countries trade disproportionately with each other, rather than with low-income, labor-abundant countries, has been used to motivate models of trade based on product differentiation rather than factor endowments. Insecurity provides a simple alternative explanation: good institutional support for trade among high-income countries lowers transactions costs, and the consequent substitution effect encourages these countries to trade disproportionately with each other. Note that our argument does not imply, counterfactually, that low-income countries should also trade disproportionately with each other.

Methodologically, this paper is closely related to the large gravity literature (Anderson, 1979; Bergstrand, 1985, 1989; Thursby & Thursby, 1987; Oguledo & MacPhee, 1994; Gould, 1994; McCallum, 1995; Frankel, Stein, & Wei, 1998; Helliwell, 1998; Feenstra, Markusen, & Rose, 2001; Baier & Bergstrand, 2001). We estimate relative import demands to deal with a common specification error suggested by Anderson's rationale for the gravity equation. We also suggest a way of estimating the price index which arises from the theoretical gravity model (Anderson, 1979). This price index is usually suppressed in empirical work because estimation has not appeared feasible.

This paper tackles just one part of the problem of resistance to trade. We do not construct a political economic model capable of explaining the emergence of good institutions. We cannot add much to the discussion of the negative effect of distance on trade (Grossman, 1998; but also see Hummels (1999)); although it is possible that traversing great distances might increase exposure to

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* Boston College and National Bureau of Economic Research, and Boston College, respectively.

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TABLE 1.—RANKINGS OF "OBSTACLES FOR DOING BUSINESS"

	Worldwide Sample
Tax regulations or high taxes	1
Corruption	2
Financing	3
Inadequate infrastructure	4
Crime and theft	5
Inflation	6
Uncertainty of cost of regulations	7
Policy instability	8
Labor regulations	9
Regulations on foreign trade	10
Safety or environmental regulations	11
Start-up regulations	12
Foreign currency regulations	13
Price controls	14
Terrorism	15

Source: Brunetti, Kisunko, and Weder (1997, p. 70).

misappropriation, our model cannot distinguish that effect from other distance-related costs. We do not consider the role of networks in reducing information costs associated with international trade, although we find such models quite plausible (Rauch & Casella, 1998; Rauch, 1999; Rauch & Trindade, 1999). Indeed, information-based models and ours overlap to the extent that information about exposure to misappropriation matters.

II. Import Demand in an Insecure World

A simple observation underlies our work: international exchange is insecure. Shipments may be hijacked. Bribes may be extorted. Contracts may not be enforced.

Our hypothesis is that insecurity constrains trade by raising the price of traded goods. In an earlier paper, we developed a model in which individuals freely allocate their labor across productive and predatory activities, endogenously determining the proportion of shipments which will be lost (Anderson & Marcouiller, 1998). We now develop a model in which the probability of loss is reflected in a price markup equivalent to a hidden tax on trade. Analogously, Anderson and Young (1999) show that imperfect contract enforcement in the importer's country is equivalent under risk neutrality to a tariff.

Our model assumes two-stage budgeting. Agents first determine the proportion of total expenditure to allocate to internationally traded goods. In a second stage, they allocate traded goods expenditure across goods. The first-stage preferences are not restricted beyond quasi-concavity. Preferences across tradable goods are CES and identical across countries. Production is specialized so that each country produces one nontraded good and one traded good. Traded goods are differentiated by country of origin.[1]

[1] Helliwell (1998, p. 10) notes other papers using this Armington assumption.

Demand in country i for the traded good produced in country j is given by

$$m_{ij} = \alpha_j \left(\frac{p_{ij}}{P_i}\right)^{-\sigma} \left(\frac{x_i}{P_i}\right) \tag{1}$$

where x_i is country i's total expenditure on traded goods,
p_{ij} is the price of j's good in i with producer prices p_{jj} normalized[2] to 1,
$P_i = [\sum_j \alpha_j p_{ij}^{1-\sigma}]^{1/(1-\sigma)}$ is the CES price index for traded goods in i, and
σ is the elasticity of substitution among traded goods.

We eliminate the product-specific parameters α_j by working with imports by i from j relative to imports by k from j:

$$\frac{m_{ij}}{m_{kj}} = \left(\frac{p_{ij}}{p_{kj}}\right)^{-\sigma} \left(\frac{P_i}{P_k}\right)^{\sigma-1} \frac{x_i}{x_k}. \tag{1'}$$

Working with relative demands addresses a specification error that plagues many empirical studies following Anderson's 1979 rationale for the gravity equation. The gravity model is derived from the import demand system by imposing the adding-up constraint that shipments to the entire world be equal to income, solving the constraint for the expenditure share for each exporter, and then substituting this exporter-specific expenditure share into the import demand equation. Following this logic, the correct specification of the gravity equation includes a highly nonlinear exporter-specific price index on the right-hand side. Focusing on imports by i and k from the same exporter j eliminates this complication, although the importer-specific price index remains.

The price of j's product in i exceeds the producer's price due to an "insecurity" markup, a transport cost, and a tariff if applicable. The deployment of a good deal of algebraic artillery[3] (see appendix) leads to the simple and plausible result that the magnitude of the insecurity markup depends on the quality of a country's institutions for the defense of trade and on factors that allow traders to use those institutions skillfully. This leads us to postulate the following relation between the price of country j's export in country i and its price in country k:

$$\frac{p_{ij}}{p_{kj}} = \left(\frac{S_i}{S_k}\right)^{\delta 1} \left(\frac{1 + b_{ij}}{1 + b_{kj}}\right)^{\delta 2} \left(\frac{1 + l_{ij}}{1 + l_{kj}}\right)^{\delta 3} \times \left(\frac{d_{ij}}{d_{kj}}\right)^{\delta 4} \left(\frac{1 + (1 - a_{ij})t_i}{1 + (1 - a_{kj})t_k}\right). \tag{2}$$

In this equation, S_i denotes the strength of the importing country's institutions for defending trade. Indicator vari-

[2] We assume here that domestic exchange is not subject to predation. This issue is addressed later and in the appendix.

[3] We thank a referee for this image.

ables b_{ij} and l_{ij} take the value of 1 if the importer and exporter share a common border or a common language, both of which encourage familiarity and may enhance the exporter's skill in using the institutions of the importing country for the defense of his or her shipments. Increasing the distance between the trading partners, d_{ij}, increases transport costs and may also expose goods to greater risk of misappropriation during shipment. The variable t_i is the importer's average ad valorem tariff rate, and a_{ij} is an indicator that takes the value 1 if the partners are associated in a free trade agreement (so that the tariff is not applied by country i to imports from country j). The tariff term in equation (2) lacks an exponent because an ad valorem tariff raises the price precisely in proportion to the tariff.

The model we present here is for simplicity one of insecurity in international trade. Difficulties in contract enforcement are most acute when multiple legal jurisdictions are involved (Rodrik, 2000). Moreover, actors within a single economy are likely to have easier access to the sort of informal enforcement mechanisms assumed by Marcouiller and Young (1995). The appendix shows that the extension to allow predation on domestic exchange is of little consequence to the interpretation of our results. If we normalize to 1 the price of j's tradable good at the factory door, its price to consumers in the home country, p_{jj}, would be affected by institutional quality just as suggested by equation (2). (See appendix.)

A country's total expenditure on traded goods, x_i, is some fraction ϕ of its total income. We model the share of total expenditure devoted to traded goods as a reduced-form function reflecting the interplay of demand and supply in general equilibrium. In the cross section, appropriate exogenous variables for the ϕ function include the country's total income, its income per capita, and its traded goods price index. The equilibrium price of the nontraded good is a reduced-form function in the same variables and is subsumed in the traded goods expenditure share function.[4] Several static structural models yield such a function.[5] Anderson (1979) rationalized this reduced form with a model of perfect competition and constant returns to scale.[6] Bergstrand (1985, 1989) developed the reduced form from a model with monopolistic competition and economies of scale.[7] Income and income per capita pick up the effect of factor endowments, possibly nonhomothetic preferences, and possible scale economies, and the traded goods price index picks up substitution between traded and nontraded goods. (Of course, the reduced-form parameters are combinations of the various structural elasticities of supply and demand.)

Imposing log linearity, we postulate as the reduced-form traded goods expenditure share

$$\phi(y_i, n_i, P_i) = \lambda y_i^{\gamma 1}(y_i/n_i)^{\gamma 2}(P_i)^{\gamma 3} \tag{3}$$

where y_i is national income and n_i is population. Then,

$$\frac{x_i}{x_k} = \frac{\phi_i y_i}{\phi_k y_k} = \frac{y_i^{(1+\gamma 1)}(y_i/n_i)^{\gamma 2}P_i^{\gamma 3}}{y_k^{(1+\gamma 1)}(y_k/n_k)^{\gamma 2}P_k^{\gamma 3}}. \tag{4}$$

Note the double role played in our model by the traded goods price index, P_i, which affects relative imports both directly through the CES demand functions and through the traded goods expenditure share.

We approximate the relative traded goods price index by a version of the Törnqvist index:

$$\ln\left(\frac{P_i}{P_k}\right) = \sum_j w_j \ln\left(\frac{p_{ij}}{p_{kj}}\right) \tag{5}$$

where w_j represents the ratio of country i's expenditure on traded good j to country i's total expenditure on all traded goods including the tradable good produced at home. It can be shown that, for a given consuming country i,

$$w_{ij} = \frac{p_{ij}m_{ij}}{\sum_{j, j\neq i} p_{ij}m_{ij}}(1 - w_{ii}).$$

We use this to construct a set of weights w_j that sum to 1 and that are identical across consumers. Ignoring the importer-specific price index, as the gravity literature sometimes has, certainly results in misspecification.[8] Our approximation is an imperfect but sensible and operational measure.

Our approach to the price index grounds theoretically a "remoteness" index like that often justified on intuitive

[4] We suppress here for simplicity any effect of insecurity on the equilibrium price of the nontraded good. See the appendix for a demonstration that this is a harmless simplification.

[5] Our empirical work explains trade in a single year, so static models are appropriate. In reality, balanced trade is rare, and the traded-goods expenditure share reflects an intertemporal margin of decision making. We ignore this margin because it is remote from the concerns of our model and seems unlikely to add to its explanatory power. Temporary trade control measures taken for balance of payments reasons will show up in the traded-goods price index.

[6] Let $e(h, p, u)$ be the expenditure function of the representative agent and let $g(h, 1, v)$ be the gross domestic product function. We have set export prices in equilibrium at unity, with landed import prices equal to the vector $p > 1$ due to transactions costs. The nontraded "home" good price is h, and the vector of factor endowments is v. Equilibrium is defined by the values of utility and nontraded goods prices consistent with balanced trade and home market clearance: $\{U(p, v), H(p, v)\} = \{u, h \mid e(h, p, u) - g(h, 1, v) = 0, e_h(h, p, u) - g_h(h, 1, v) = 0\}$. Here we assume that tariff revenue is lost to rent-seeking or government corruption; hence, expenditure equals production income. We use Shephard's and Hotelling's lemmas in defining market clearance. The reduced-form traded goods expenditure share is defined as $\phi(p, v) \equiv p \cdot e_p[H(p, v), p, U(p, v)]/e[\cdot]$. We further simplify by using a CES aggregator of traded-goods prices $P(p)$.

[7] Oguledo and MacPhee (1994) follow Bergstrand.

[8] The previously cited papers by Bergstrand and by Gould also address this issue.

grounds because $\sum_j w_j \ln(d_{ij}/d_{kj})$ is an element of $\sum_j w_j \ln(p_{ij}/p_{kj})$. Analogous intuitions can be offered for other terms of the price index. For example, bilateral trade is expected to be greater when the partners speak a common language than when they do not. However, the "common language" effect on bilateral trade might well be greater for two countries that speak Dutch, a relatively uncommon language, than for two English-speaking countries with many potential same-language partners. Because $\sum_j w_j \ln((1 + l_{ij})/(1 + l_{kj}))$ is an element of $\sum_j w_j \ln(p_{ij}/p_{kj})$, our treatment of the price index ought to pick up any such second-order effect. The strength of our approach is that we are able to define appropriate weights and a common treatment of the set of such intuitively plausible effects.

Substituting equations (2) and (4) into equation (1′) and taking logs,

$$\begin{aligned}\ln\frac{m_{ij}}{m_{kj}} &= (1 + \gamma_1)\ln\left(\frac{y_i}{y_k}\right) + \gamma_2 \ln\left(\frac{y_i/n_i}{y_k/n_k}\right) \\ &+ (\sigma - 1 + \gamma_3)\ln\left(\frac{P_i}{P_k}\right) - \sigma\delta_1 \ln\left(\frac{S_i}{S_k}\right) \\ &- \sigma\delta_2 \ln\left(\frac{1 + b_{ij}}{1 + b_{kj}}\right) - \sigma\delta_3 \ln\left(\frac{1 + l_{ij}}{1 + l_{kj}}\right) \\ &- \sigma\delta_4 \ln\left(\frac{d_{ij}}{d_{kj}}\right) - \sigma \ln\left(\frac{1 + (1 - a_{ij})t_i}{1 + (1 - a_{kj})t_k}\right).\end{aligned} \tag{6}$$

Using equations (5) and (2) again,

$$\begin{aligned}\ln\left(\frac{P_i}{P_k}\right) &= \delta_1 \sum_j w_j \ln\left(\frac{S_i}{S_k}\right) + \delta_2 \sum_j w_j \ln\left(\frac{1 + b_{ij}}{1 + b_{kj}}\right) \\ &+ \delta_3 \sum_j w_j \ln\left(\frac{1 + l_{ij}}{1 + l_{kj}}\right) + \delta_4 \sum_j w_j \ln\left(\frac{d_{ij}}{d_{kj}}\right) \\ &+ \sum_j w_j \left(\frac{1 + (1 - a_{ij})t_i}{1 + (1 - a_{kj})t_k}\right).\end{aligned} \tag{7}$$

Two simplifications come into play when we substitute equation (7) into (6). The ratio S_i/S_k does not vary across exporters j. Therefore,

$$\sum_j w_j \ln\left(\frac{S_i}{S_k}\right) = \ln\left(\frac{S_i}{S_k}\right).$$

Similarly, because relatively few of our bilateral trade flows involve free trade agreements causing average tariffs to vary across exporters, we also collapse the "weighted sum" of the relative tariffs into the original tariff term. This leaves us with the following model of imports by country i from j relative to imports by the base country k from the same exporter j:

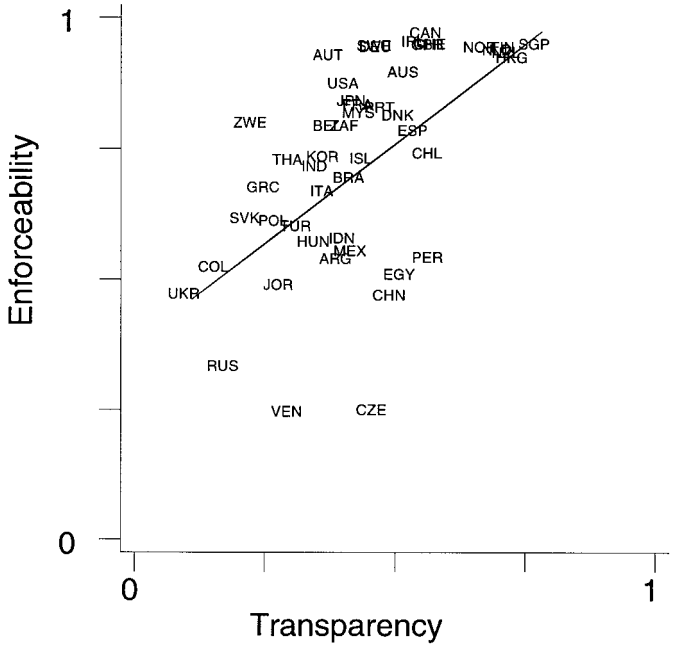
$$\begin{aligned}\ln\frac{m_{ij}}{m_{kj}} &= (1 + \gamma_1)\ln\left(\frac{y_i}{y_k}\right) + \gamma_2 \ln\left(\frac{y_i/n_i}{y_k/n_k}\right) \\ &+ (\gamma_3 - 1)\delta_1 \ln\left(\frac{S_i}{S_k}\right) - \sigma\delta_2 \ln\left(\frac{1 + b_{ij}}{1 + b_{kj}}\right) \\ &- \sigma\delta_3 \ln\left(\frac{1 + l_{ij}}{1 + l_{kj}}\right) - \sigma\delta_4 \ln\left(\frac{d_{ij}}{d_{kj}}\right) \\ &+ (\gamma_3 - 1)\ln\left(\frac{1 + (1 - a_{ij})t_i}{1 + (1 - a_{kj})t_k}\right) \\ &+ (\sigma - 1 + \gamma_3)\left[\delta_2 \sum_j w_j \ln\left(\frac{1 + b_{ij}}{1 + b_{kj}}\right)\right. \\ &\left. + \delta_3 \sum_j w_j \ln\left(\frac{1 + l_{ij}}{1 + l_{kj}}\right) + \delta_4 \sum_j w_j \ln\left(\frac{d_{ij}}{d_{kj}}\right)\right].\end{aligned} \tag{8}$$

Note that the variable of primary interest to us—security—potentially affects prices and relative imports through three channels: the substitution effect within the group of traded goods, $(p_{ij}/P_i)^{-\sigma}$; the general equilibrium substitution effect between traded and nontraded goods, $\partial\phi/\partial P_i$; and the real income effect, $\partial(x_i/P_i)/\partial P_i$. Country i's security, S_i, does not affect the relative prices it faces within the set of traded goods, p_{ij}/P_i, so the first of the substitution effects is nil. The reduced-form coefficient on the institutional quality index, $(\gamma_3 - 1)\delta_1$, is the product of δ_1, which captures the elasticity of the price index P_i with respect to S_i, and $\gamma_3 - 1$, which captures the net elasticity of trade with respect to P_i. The term $\gamma_3 - 1$ is itself the sum of the real income elasticity, -1, and the general equilibrium substitution elasticity, γ_3 in equation (4).[9]

All the major elements of our model are now in place. We have modeled a world in which traded goods are differentiated by country of origin. Differences across importers in demand for a single good have three sources: (i) differences in the price markups associated with insecurity, distance, and tariffs, (ii) differences in total expenditure, and (iii) differences in the division of expenditure between traded and nontraded goods.

[9] The appendix shows that a reduced-form coefficient for the institutional index with a more complex interpretation emerges from a model in which domestic exchange is also insecure. Nothing in what follows hangs on the particular "structural" interpretation of the institutional index coefficient.

FIGURE 1.—TRANSPARENCY AND ENFORCEABILITY SCORES



III. Data

The security of exchange depends on the strength of two sets of institutions: those that support transactions among private firms and those that support transactions between the private sector and the state. We rely on data provided by the World Economic Forum (WEF) to measure the quality of both sets of institutions. The measures are drawn from the WEF 1997 Executive Survey, which was completed by more than 3,000 participants distributed across 58 countries (World Economic Forum 1997, p. 85). Participants in the WEF survey were asked to assign a score ranging from 1 (strongly disagree) to 7 (strongly agree) to each of the following statements:

- Government economic policies are impartial and transparent.
- The legal system in your country is effective in enforcing commercial contracts.

We rescale the mean response for each country to run between 0 and 1. Figure 1 plots each country's score on "enforceability" against its score on "transparency."

Admittedly, these are noisy signals of institutional strength. Expectations differ across countries, so that what counts as "effective" enforcement or "impartial" policy in Russia may differ from what would be similarly classified in Singapore. The respondents to the survey form a selected group; even if they were randomly selected within a country, they would still represent only those firms that had chosen not to relocate or to shut down. Moreover, the WEF provides only the mean response for each country; we lack information about within-country variation in responses.

As a check on the robustness of our results, we also use a complementary "composite security" index formed from responses to the following eleven statements:

- Government economic policies are impartial and transparent.
- Government regulations are precise and fully enforced.
- Tax evasion is minimal.
- Irregular additional payments are not common in business and official transactions.
- The legal system is effective in enforcing commercial contracts.
- Agreements and contracts with the government are not often modified due to budget cutbacks, changes in government, or changes in government priorities.
- Private businesses can readily file lawsuits at independent and impartial courts if there is a breach of trust on the part of the government.
- New governments in your country honor the commitments and obligations of previous regimes.
- Citizens of your country are willing to adjudicate disputes rather than depending on physical force or illegal means.
- Your country's police are effective in safeguarding personal security so that this is not an important consideration in business activity.
- Organized crime does not impose significant costs on business in your country.

Respondents were asked to rank agreement with each statement on a scale from 1 to 7. We took the country mean responses to each of the eleven statements, rescaled them to run between 0 and 1, took logs, and then used factor analysis to extract the first principal factor.[10] Table 2 shows the composite security score for each of the countries.

It has been suggested that our indices of institutional quality may act as proxies for traditional barriers to trade. However, tariff barriers and trade preferences enter our model explicitly, and the correlation between institutional measures and nontariff barrier (NTB) coverage ratios is low for those countries for which we have data.

[10] Scoring a new variable using the loadings for the first principal factor generates a variable with mean of 0 and a standard deviation of 1. One can't take the log of such a variable. Instead, we have taken the logs of the mean responses to the eleven questions before extracting the first principal factor and scoring the new variable. Because the factor extraction is based on the correlation matrix, the composite indicator is invariant to the choice of base country k.

TABLE 2.—COMPOSITE SECURITY SCORES

Importer	Score
Russia	−2.614
Ukraine	−2.377
Venezuela	−2.218
Colombia	−2.098
Greece	−1.195
Poland	−0.858
Thailand	−0.796
Jordan	−0.794
Hungary	−0.791
Mexico	−0.749
South Africa	−0.602
Argentina	−0.579
Turkey	−0.539
Slovak Republic	−0.524
Brazil	−0.521
Czech Republic	−0.452
Italy	−0.362
Indonesia	−0.284
India	−0.264
Zimbabwe	−0.240
Peru	−0.235
Korea	−0.217
China	−0.184
Belgium-Luxembourg	0.055
Egypt	0.227
Spain	0.382
Portugal	0.391
Iceland	0.451
Malaysia	0.499
Japan	0.562
United States	0.651
Chile	0.680
France	0.689
Australia	0.704
Sweden	0.779
Austria	0.807
Denmark	0.857
Ireland	0.864
Germany	0.931
New Zealand	0.997
United Kingdom	1.034
Netherlands	1.036
Canada	1.050
China: Hong Kong	1.134
Norway	1.142
Switzerland	1.159
Finland	1.173
Singapore	1.241

The correlation coefficient between the NTB coverage ratio and our index of transparency is −0.32, the correlation with our index of enforceability is −0.14, and the correlation with our composite security index is −0.15.[11] The signs are those that one might expect from a political economy perspective, but the magnitudes of the correlations are small.

Our data on 1996 bilateral import expenditures are taken from the *Direction of Trade Statistics* (DOTS) published by the International Monetary Fund. Most of the DOTS import data are reported c.i.f.; some appear only f.o.b. To avoid as much as possible ad hoc adjustments to the data, we use the reported c.i.f. figures, adjusting the few f.o.b. figures upward by a factor based on the ratio between the country's total reported c.i.f. imports from the rest of the world and the world's total reported exports to that country.[12]

Data on 1996 population and GDP in current dollars are taken from the World Bank's *World Development Indicators* (WDI). We calculate distance from capital city to capital city on the basis of geographical coordinates listed by Fitzpatrick and Modlin (1986); of course, the distance from Washington to Ottawa only roughly captures the average distance traversed by shipments from the United States to Canada. David Tarr and Francis Ng of the World Bank graciously provided us with unweighted average tariff data; this series is far more complete than the data on import duties as a percentage of import expenditures reported in the WDI.[13] We composed dummy variables to capture sharing a common border, a common language, or common membership in ASEAN, the EU, MERCOSUR, or NAFTA.

We have complete data on these variables for a total of 2,182 import flows distributed across 48 importing countries. Table 3 shows the importing countries in our data set and the number of import flows that we observe for each. For an additional 24 bilateral pairs (1% of the total), no imports were reported.[14]

IV. Estimation and Results

Our structural model leads to a simple result. Relative import demand is a function of the relative levels of income, income per capita, institutional quality, a shared border, a shared language, distance, tariffs if applicable, and the overall traded goods price index. Estimation of the model supports three contentions:

(i) By lowering transactions costs, institutional support for secure exchange significantly raises international trade volume.

(ii) Excluding institutional variables obscures a negative relation between income per capita and the share of income spent on traded goods, other things equal.

(iii) Institutional differences can generate "a disproportionately high volume of trade among high-income countries," a pattern "which happens to accord well with trade patterns in the real world" (Deardorff, 1998, p. 16).

[11] The nontariff barrier coverage ratios are taken from the WEF's *Global Competitiveness Report 1997*, p. 223. They are available for only 37 of our 48 countries.

[12] Estimates of our model over interpolated f.o.b. import flows, applying analogous factors of adjustment to deflate the c.i.f. import values to approximate f.o.b. equivalents, are available on request.

[13] Even so, not every country has tariff data available for 1996. We have used 1996 data where available, but data from 1997, 1995, or 1994 where necessary.

[14] That is, the country pair appears in the DOTS data matrix but the trade volume is given as ".".

We begin by estimating equation 3.1, a stochastic form of the model developed in section II (equation (8)):

$$\begin{aligned}\ln\frac{m_{ij}}{m_{kj}} = {} & \beta_0 + \beta_1 \ln\left(\frac{y_i}{y_k}\right) + \beta_2 \ln\left(\frac{y_i/n_i}{y_k/n_k}\right) + \beta_3 \ln\left(\frac{S_i}{S_k}\right) \\ & + \beta_4 \ln\left(\frac{1+b_{ij}}{1+b_{kj}}\right) + \beta_5 \ln\left(\frac{1+l_{ij}}{1+l_{kj}}\right) \\ & + \beta_6 \ln\left(\frac{d_{ij}}{d_{kj}}\right) + \beta_7 \ln\left(\frac{1+(1-a_{ij})t_i}{1+(1-a_{kj})t_k}\right) \\ & + \beta_8 \sum_j w_j \ln\left(\frac{1+b_{ij}}{1+b_{kj}}\right) \\ & + \beta_9 \sum_j w_j \ln\left(\frac{1+l_{ij}}{1+l_{kj}}\right) \\ & + \beta_{10} \sum_j w_j \ln\left(\frac{d_{ij}}{d_{kj}}\right) + \upsilon_i + \epsilon_{ij}\end{aligned} \tag{9}$$

The stochastic term includes two elements. The first, υ_i, captures any disturbance that systematically affects all of country i's imports relative to those of the base country k. The second element is specific to imports by i from exporter j, ϵ_{ij}. The base country k is held constant. We estimate the model using Stata's White correction for possible heteroskedasticity, with clustering by importer. The standard errors estimated with clustering are approximately twice those derived from a simple White correction, indicating significant positive correlation of residuals within importer-specific clusters.[15]

Table 4 reports summary statistics for the import, GDP, GDP per capita, transparency, enforceability, composite security, common border, common language, distance and tariff ratios (as defined previously), using the United States as a convenient base country, k.[16]

Table 5 reports the results of estimating equation (9) under various restrictions. The last two columns show estimates from tobit models, with the 24 unreported import flows taken to be zero (and left-censored).[17] All six columns show robust standard errors with clustering by importer.

The first of our conclusions is that, by lowering transactions costs, institutional support for secure exchange significantly increases trade. The composite security index has a positive and highly significant effect on imports (Reg 4). The transparency and enforceability indices have independent positive effects (Reg 3), highly significant in the case of transparency and marginally so in the case of enforceability.

A few examples shed light on the magnitude of the effects implied by the point estimates. Transparency is rated roughly 10% higher in France than in Argentina. Interpreting the estimated coefficient on transparency as a reduced-form elasticity, this difference implies roughly 5% higher imports into France than into Argentina, other things equal.

If the seven Latin American countries in our sample (Argentina, Brazil, Chile, Colombia, Mexico, Peru, and

TABLE 3.—IMPORTERS IN THE DATA SET

Importer	No. Obs.
Argentina	46
Australia	47
Austria	46
Belgium-Luxembourg	47
Brazil	47
Canada	47
Chile	36
China	47
China: Hong Kong	47
Colombia	46
Czech Republic	47
Denmark	47
Egypt	47
Finland	47
France	47
Germany	47
Greece	46
Hungary	47
Iceland	42
India	47
Indonesia	46
Ireland	47
Italy	47
Japan	47
Jordan	42
Korea	33
Malaysia	46
Mexico	38
Netherlands	47
New Zealand	47
Norway	46
Peru	45
Poland	47
Portugal	47
Russia	47
Singapore	44
Slovak Republic	47
South Africa	47
Spain	47
Sweden	47
Switzerland	46
Thailand	43
Turkey	47
Ukraine	41
United Kingdom	47
United States	47
Venezuela	45
Zimbabwe	42
Total	2182

[15] Parameter estimates will be biased if the covariance of GDP and υ_i is not zero. In our cross-sectional context, one reasonable approach to screen out the effect of a contemporaneous shock to both y_i and m_{ij} is to use lagged GDP as an instrument for current GDP. Using this procedure gave us parameter estimates well within one standard error of the ones reported here, and it strengthened, if anything, the security and home bias effects. Results available on request.

[16] Similar results were found when Brazil and China were used as alternative base countries and also when all U.S. trade was dropped from the sample. Results are available on request.

[17] With an elasticity of substitution among traded goods that exceeds 1, high transactions costs can eliminate trade in some bilateral pairings. In this case, the log of the import ratio, ln(0), was assigned a value 0.1 below the log of the lowest positive import ratio in the data set.

TABLE 4.—RATIOS WITH UNITED STATES AS BASE COUNTRY

Ratio: U.S. as Base	Number Observations	Mean	Standard Deviation
Import ratio	2135	0.281	0.977
GDP ratio	2135	0.079	0.173
GDP per capita ratio	2135	0.520	0.441
Transparency ratio	2135	1.085	0.370
Enforceability ratio	2135	0.833	0.226
Composite security ratio	2135	0.012	0.981
Common border ratio	2135	1.026	0.238
Common language ratio	2135	0.948	0.263
Distance ratio	2135	1.204	1.848
Tariff ratio	2135	1.035	0.068

Venezuela) were to enjoy the same transparency and enforceability scores as the mean scores of the members of the European Union, predicted Latin American import volumes would rise 30%.[18] This increase is of roughly the same magnitude as the 35% increase that could be expected from lowering Latin American tariffs to the levels applied by the United States, holding other things equal. A much larger (51%) increase in average Latin American GDP would be necessary to generate a comparable increase in imports.

As can be seen from equation (8), these "thought experiments" involve several distinct effects. The calculations multiply the effect of insecurity on the price markup (δ_1) by (i) the reduced-form substitution effect on the division of expenditure between traded goods and the nontraded good (γ_3), and (ii) the real income effect of the price change (-1). In the notation of section II, the net effect of the security term is $(\gamma_3 - 1)\delta_1$.

The coefficient on the tariff variable in table 5 represents an estimate of $(\gamma_3 - 1)$. When the institutional variables are included, we estimate the reduced-form elasticity of import demand with respect to tariffs as roughly -4.8, an estimate remarkably similar to the elasticity estimated by Baier and Bergstrand (2001). The implied estimate of $\hat{\gamma}_3 = -3.8$ suggests that the substitution effect of insecurity on the trade share outweighs the real income effect, which is equal to -1.

Comparison of equations (8) and (9) implies that $\hat{\beta}_4/\hat{\beta}_8$, $\hat{\beta}_5/\hat{\beta}_9$, and $\hat{\beta}_5/\hat{\beta}_{10}$ should all be equal. *F*-tests on the estimated coefficients for the full model (Reg 3 and Reg 4) do not reject that hypothesis.[19] Note that the positive sign on the weighted distance variable makes sense in terms of our model if the elasticity of substitution among traded goods is high. Assuming that p_{ij} rises with the distance between i and j, we expect δ_4 to be positive. The positive estimated coefficient on the weighted distance term then requires that $\gamma_3 + \sigma - 1 > 0$. Using the estimate of $\hat{\gamma}_3 = -3.8$ from the estimated coefficient on the tariff ratio gives us a lower bound on the elasticity of substitution among traded goods: σ must be above 4.8.

One can push the structural interpretation of the coefficients to its limit by estimating equation (8) directly using maximum likelihood techniques. The estimated parameters and the associated robust standard errors (with clustering by importer, as always) are displayed in table 6. As expected, the estimated elasticity of substitution among the traded goods, $\hat{\sigma}$, is high. Higher security lowers the price p_{ij}, as captured by the parameters $\hat{\delta}_{1a}$ and $\hat{\delta}_{1b}$ when transparency and enforceability are used and by $\hat{\delta}_1$ when the composite security score is used. By lowering p_{ij} for all j, enhanced security also lowers P_i, with a positive effect on the traded goods expenditure share (through $\hat{\gamma}_3$). We do not wish to emphasize these maximum likelihood results, however, because they come from a model that is very highly constrained. The unconstrained regression results of table 5 lead us to question the empirical importance of the weighted language and border variables, to which the constrained maximum likelihood model gives a good deal of influence as it forces the cross-term parameter restrictions to be met.

Our second major proposition is that higher income per capita reduces the share of expenditure devoted to traded goods, all else equal. The estimated coefficients on the log of per capita GDP in table 5 are our estimates of γ_2, the reduced-form elasticity of the traded goods expenditure share with respect to income per capita. As shown in table 5, the estimates of this elasticity change sign as tariff and security variables are added to the regression. Inappropriate exclusion of the tariff and security variables leads to the result shown in the first column, with a "significantly" positive coefficient on GDP per capita.[20] The coefficient becomes insignificantly different from zero when the tariff term is added. Including enforceability and transparency drives the coefficient into the negative range. Other things equal, doubling per capita income reduces the share of total expenditure devoted to traded goods by 20%.

Our anti-trade bias result, which bears a resemblance to earlier work by Hunter and Markusen (1988), is what one expects if the proportion of expenditure devoted to nontraded services rises as income per capita rises.[21] Our result stands in contrast to recent work that has found expenditure shares to be invariant to income (Davis et al., 1997, pp. 433–436). It stands in even sharper contrast to work that finds a positive effect of income per capita on trade (Frankel

[18] In response to the improvement in the transparency score, the log of relative imports would be expected to rise by 0.095, the product of the regression coefficient, 0.530, times the increase in log transparency from ln(0.99) to ln(1.19). The increase in log enforceability from ln(0.62) to ln(0.98) times its regression coefficient, 0.385, would raise the log of relative imports by 0.174. Exponentiating, the total increase of 0.269 in log imports is equivalent to multiplying import levels by 1.30, which we interpret as a rise of 30%. A similar procedure underlies the other comparative statics calculations.

[19] Using Reg 3, the *F*-statistic for the joint hypothesis that $\hat{\beta}_4/\hat{\beta}_8 = \hat{\beta}_6/\hat{\beta}_{10}$ and $\hat{\beta}_5/\hat{\beta}_9 = \hat{\beta}_6/\hat{\beta}_{10}$ is $F(2, 47) = 1.08$. Using Reg 4, the statistic is $F(2, 47) = 1.90$.

[20] Of course, it is our contention that Reg 1 suffers from omitted-variables bias.

[21] The reduced form permits no structural interpretation. A natural candidate is departures from homotheticity, but our result can be found in a model with homothetic preferences and constant returns technology.

TABLE 5.—RELATIVE IMPORT DEMAND, UNITED STATES AS THE BASE

Variable	Reg 1	Reg 2	Reg 3	Reg 4	Tobit 1	Tobit 2
Log GDP ratio	0.837	0.855	0.860	0.866	0.907	0.911
	(0.045)	(0.042)	(0.037)	(0.038)	(0.037)	(0.040)
Log GDP per capita ratio	0.141	0.018	−0.206	−0.191	−0.244	−0.215
	(0.058)	(0.094)	(0.105)	(0.122)	(0.108)	(0.125)
Log transparency ratio	•	•	0.530	•	0.620	•
	•	•	(0.169)	•	(0.173)	•
Log enforceability ratio	•	•	0.385	•	0.307	•
	•	•	(0.199)	•	(0.196)	•
Relative composite security	•	•	•	0.285	•	0.279
	•	•	•	(0.073)	•	(0.081)
Log common border ratio	0.908	0.794	0.753	0.747	0.668	0.665
	(0.140)	(0.155)	(0.160)	(0.163)	(0.184)	(0.186)
Log common language ratio	0.314	0.327	0.331	0.336	0.349	0.358
	(0.081)	(0.080)	(0.082)	(0.082)	(0.109)	(0.109)
Log distance ratio	−1.134	−1.109	−1.097	−1.095	−1.134	−1.133
	(0.054)	(0.058)	(0.056)	(0.056)	(0.055)	(0.056)
Log adjusted tariff ratio	•	−2.973	−4.753	−4.814	−4.773	−4.699
	•	(1.992)	(2.146)	(2.343)	(2.126)	(2.327)
Weighted log border ratio	−1.807	−1.654	−1.092	−1.391	−0.934	−1.169
	(1.474)	(1.378)	(1.332)	(1.364)	(1.363)	(1.426)
Weighted log language ratio	1.390	1.438	−0.001	−0.119	0.809	0.637
	(1.639)	(1.486)	(1.448)	(1.363)	(1.417)	(1.433)
Weighted log distance ratio	0.420	0.424	0.382	0.451	0.300	0.386
	(0.164)	(0.160)	(0.137)	(0.130)	(0.126)	(0.129)
Constant	0.055	0.076	−0.169	−0.184	−0.142	−0.131
	(0.158)	(0.146)	(0.135)	(0.147)	(0.121)	(0.144)
Number observations	2135	2135	2135	2135	2159	2159
R^2	.69	.69	.70	.70		
Log likelihood					−3859	−3865

Robust standard error (Huber/White) with clustering by importer given in parentheses. Imports are c.i.f., as reported in *Direction of Trade Statistics.* Results using interpolated f.o.b. figures are available on request.

et al., 1998, p. 96), an effect which we reproduce only by excluding the tariff and insecurity variables.

Econometrically, these changes in sign are driven by correlation between GDP per capita and the omitted variables. The correlation coefficient between GDP per capita and the tariff ratio is −0.62. When the tariff ratio is dropped from the regression, part of the positive effect of lower tariffs on trade is misread as a positive effect of higher income per capita on trade. The correlation between GDP per capita and the transparency ratio is 0.55, and its correlation with the enforcement ratio is 0.73.[22] When the tariff and institutional variables are dropped from the regression, part of the positive effect of security on trade is misattributed to income per capita. Including the theoretically appropriate regressors reveals that GDP per capita actually has a negative effect on the traded goods expenditure share. Although the effect is imprecisely estimated, a rise in income per capita seems likely to lower the share of a country's total income that is spent on traded goods, other things equal.[23] Of course, to the extent that countries with higher income per capita also enjoy better institutions for the defense of trade, the trade-restricting effect of income per capita could be offset by the trade-enhancing effect of lower traded-goods prices.

Our third main contention is that institutional differences can generate "a disproportionately high volume of trade among high-income countries," a pattern "which happens to accord well with trade patterns in the real world" (Deardorff, 1998, p. 16). Why should high-income countries skew their trade toward imports from other high-income countries in spite of the presumed similarity of factor endowment? And what answer to the first question can be consistent with the stylized fact that low-income countries do not rely disproportionately on imports from other low-income countries?

Several solutions to the puzzle have been proposed (notably Markusen (1986)). We offer an explanation based on the price markup associated with insecure trade. Effective institutions in the importing country lower transactions costs, lower the prices of traded goods, and raise imports, holding constant the characteristics of the exporting country. The absolute price markup, however, would also depend on the quality of institutions in the exporting country. Our empirical results confirm that low security in country i lowers m_{ij}/m_{kj}; the price markup model also implies that both m_{ij} and m_{kj} will be low when the security of country j is low. We cannot estimate this second effect, because the impact of the exporter's security and the effect of the expenditure share α_j are not separately identified. The prediction of the model, however, clearly coincides with the observed pattern of trade. Trade among high-income coun-

[22] This correlation is given in the data, but it does not imply that income per capita and institutional quality are necessarily linked, nor does it invalidate the "thought experiment" previously reported in which institutions were improved without a corresponding increase in income per capita.

[23] In contrast, the anti-trade bias effect of size, the negative effect on trade shares of country size, measured by GDP, is unaffected by inclusion of security variables. Size is only weakly correlated with security.

TABLE 6.—MAXIMUM LIKELIHOOD PARAMETER ESTIMATES CONSTRAINED MODEL (EQUATION (8))

Parameter	Model 1	Model 2
$\gamma_1 + 1$	0.854	0.859
	(0.036)	(0.037)
γ_2	−0.211	−0.194
	(0.107)	(0.122)
$\gamma_3 - 1$	−4.801	−4.850
	(2.193)	(2.381)
δ_{1a}	−0.110	•
	(0.037)	•
δ_{1b}	−0.084	•
	(0.053)	•
δ_1	•	−0.059
	•	(0.020)
δ_2	−0.094	−0.082
	(0.058)	(0.055)
δ_3	−0.042	−0.037
	(0.024)	(0.023)
δ_4	0.139	0.120
	(0.068)	(0.066)
σ	7.915	9.125
	(3.815)	(4.962)
Constant	−0.194	−0.211
	(0.127)	(0.136)
Observations	2135	2135
Log likelihood	−3410	−3414

Model 1 includes the transparency and enforceability indices. Model 2 substitutes the composite security factor. Huber/White robust standard errors with clustering by importer in parentheses.

tries with high-quality institutions ought to be high because the transactions costs associated with insecurity are low; transactions costs impose a double disadvantage on trade among low-income, low-security countries. This solves a problem alluded to in Deardorff's (1998, p. 16) informal exposition of an explanation based on identical but nonhomothetic preferences. Our story implies disproportionate trade among consumers of the "high-income" good, but it does not imply counterfactually a similarly disproportionate amount of trade among low-income consumers.

V. Conclusion

Abundant evidence suggests that transactions costs associated with insecure exchange significantly impede international trade. Predation by thieves or by corrupt officials generates a price markup equivalent to a hidden tax or tariff. These price markups significantly constrain trade where legal systems poorly enforce commercial contracts and where economic policy lacks transparency and impartiality.

This paper builds a structural model of import demand in an insecure world and estimates the model using data collected by the World Economic Forum. We find that a 10% rise in a country's index of transparency and impartiality leads to a 5% increase in its import volumes, other things equal. Costs associated with institutional weakness beg for serious consideration as we try to explain why countries trade so little.

We also find that the share of total expenditure devoted to traded goods declines as income per capita rises, other things equal. This result stands in contrast to recent papers that fail to reject homotheticity and calls into question the frequent use of homothetic preferences in trade models.

Finally, the paper suggests an explanation for the stylized fact that high-income, capital-abundant countries trade disproportionately with each other. These countries also enjoy strong institutions for the support of trade. Because the traded-goods price markup depends on the degree of insecurity in the exporting and the importing countries, trade among rich countries will be relatively unhampered by security-related transactions costs, whereas trade among poor countries will be doubly disadvantaged.

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APPENDIX

Modeling the Security of Trade

This appendix provides a multicountry generalization of the predation model found in Anderson and Marcouiller (1998), illustrating the logic underlying equation (2). The contract enforcement model of Anderson and Young (1999) provides an alternative route to the conclusion that insecurity leads to a markup in price.

In our model, thieves congregate at borders. We think of these borders as international boundaries, but they could also be the perimeters of the markets in which exchange takes place. A complete transaction involves crossing two borders: the exporter must evade predation on the way out of his country, and the importer must evade predation when bringing the good into hers. Any shipment defended by less than the usual and customary measures would be identifiable as easy prey, attacked, and lost. Under these conditions, all shippers will take normal defensive measures and thieves will attack randomly.

The odds of predator success in the game of hide-and-seek at the border of j are plausibly given by $\theta L_j^B/L_j^D$, a function of the supply of bandits along j's border (L_j^B), the resources devoted to defending shipments at the border (L_j^D), and an exogenous technological parameter (θ). Then, the probability that an exporter of good j will successfully negotiate the border of his own country is given by the evasion success function:[24]

$$\pi_j = \frac{1}{1 + \theta\dfrac{L_j^B}{L_j^D}}. \tag{A.1}$$

The ability to diversify risk makes $(1 - \pi_j)$ equivalent, from the shippers' point of view, to a proportional insecurity tax on the value of every shipment. This tax is bounded on the unit interval, increasing in bandit labor and decreasing in defensive resources.

Anderson and Marcouiller (1998) fully endogenize the allocation of labor to defense in a two-country model. In this paper, we treat defensive arrangements L_j^D as given, for the sake of tractability in a multicountry model. We also assume the world's total supply of thieves to be exogenous: $L^B = \Sigma_j L_j^B$. However, each bandit is free to choose the location (the particular border) at which to attack shipments. The endogenous allocation of bandits to different borders then determines π_j.

Bandits maximize expected loot $\Sigma_j(1 - \pi_j(L_j^B, L_j^D, \theta))v_j$, where v_j is the volume of trade flowing through the border of country j. (The assumption that uncoordinated bandits take trade volumes as given greatly simplifies this problem.) Solving the first-order conditions gives the allocation of bandit labor to each border:

$$L_j^B = \frac{\pi_j(1 - \pi_j)v_j}{\Sigma_j\pi_j(1 - \pi_j)v_j} L^B. \tag{A.2}$$

A bit of algebra produces the reduced-form solution for π_j:

$$\pi_j = \left[\frac{L_j^D}{v_j}\right]^{1/2} \frac{\Sigma_j w_j (L_j^D/v_j)^{1/2}}{\dfrac{\theta L^B}{\Sigma_j v_j} + \Sigma_j\, w_j(L_j^D/v_j)}, \tag{A.3}$$

where w_j is country j's share of total world trade. Let $S_j^* \equiv (L_j^D/v_j)^{1/2}$ denote the strength of a country's institutions for the defense of trade. Then,

$$\pi_j = S_j^* Z, \tag{A.4}$$

where Z is simply the constant term on the right in equation (A.3). We proxy S^* with a log linear function of observable characteristics S.

The relationship between insecurity and the price markup is as follows. For simplicity, we disregard other sources of transactions cost, including tariffs. The probability of success on j's sales to i is $\pi_{ij} = \pi_i\pi_j$. Risk-neutral traders will arbitrage until $\pi_{ij}p_{ij} = p_j = 1$, where the exporter's price is set equal to 1 by choice of units. In the text, we deal with relative imports and relative prices. Note that $\pi_{ij}p_{ij} = \pi_i\pi_j p_{ij} = 1$ $\forall\ i$ implies that $p_{ij}/p_{kj} = \pi_k/\pi_i = (S_i/S_k)^\delta$ with $\delta < 0$, using equation (A.4) and the proxy S (treated as a scalar). This is the story underlying the incorporation of institutional quality indicators into equation (2).

Now extend our model to include the potential effect of insecurity on domestic transactions. Consider a world in which goods are purchased by specialized brokers at the factory door, then taken to a centralized market. Exchange takes place in the market, and then another shipper takes the goods to the final consumer. Thieves allocate themselves across market portals to maximize expected loot, different routes are defended with different resources, and the probability of success is determined as before.

The exportable good is also sold at home, with domestic price $p_{jj} > 1$ to consumers. The effect of domestic insecurity on the price index of tradable goods acts through the domestically produced tradable good exactly as it does through traded goods imported from abroad, so no alteration is needed in the model of the text.

Domestic transactions in the nontradable good are affected by insecurity in a more complex way. The traded goods expenditure share must be modified to $\phi_i = \phi(Y_i, N_i, P_i, S_i)$. Assume that home-market insecurity is solely in the final goods market because a full model of domestic insecurity would include intermediate goods and primary factor markets as well, taking us far beyond the paper's scope. The producer's price of the nontradable good is h and the consumer's price is h/π. Let $e(h/\pi, P, u)$ be the expenditure function of the representative agent, and let $g(h, 1, v)$ be the gross domestic product function. Here, the price index of traded goods facing consumers is P and the vector of factor endowments is v. Equilibrium is defined by the values of utility and home goods price consistent with balanced trade and home market clearance:

$$\{U(P, \pi, v), H(P, \pi, v)\} = \{u, h | e(h/\pi, P, u) - g(h, 1, v) = 0,$$
$$e_h(h/\pi, P, u) - g_h(h, 1, v) = 0\}.$$

We use Shephard's and Hotelling's lemmas in defining market clearance. The reduced-form traded goods expenditure share is defined as

$$\phi(P, \pi, v) \equiv P \cdot e_P[H(P, \pi, v), P, U(P, \pi, v)]/e[\cdot].$$

Finally, replace π with the proxy S and replace v with (Y, N). Imposing log linearity, essentially the same regression model emerges. The difference is that the effect of insecurity on trade now includes its effect directly on the reduced-form trade expenditure share, not just its effect via the price index.

The interpretation of our results in the text is not much affected by this extension to domestic insecurity. The reduced-form trade expenditure share is lowered by insecurity insofar as it raises P, the price index of tradable goods, and is raised by insecurity insofar as it raises h/π, the consumer price of the nontraded "home" good. Our results show that trade-reducing effects predominate.

[24] The same function has been used to model the outcome of contests between a single predator and his prey by Grossman and Kim (1995). Despite the formal similarity, our model of group interaction is distinct. We are concerned with the probability of an encounter between predator and prey, whereas Grossman and Kim are concerned about the outcome of the contest given an encounter.