Essays on Conflict, Corruption, and International Trade Politics

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ESSAYS ON CONFLICT, CORRUPTION, AND INTERNATIONAL TRADE POLITICS

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ABSTRACT

This dissertation is a collection of three essays which examine issues at the intersection of international economics, political economy, and macroeconomics. A common theme which emerges in the subsequent chapters is a reliance on intuitive models of economies populated by rational agents engaging in both political and economic decisionmaking. Each chapter also presents empirical evidence using aggregate data to highlight new angles on issues related to macroeconomic development policy.

Concurrent cross-country political change, such as the recent "Arab spring" revolutions in the Middle East, the experience of South American military dictatorships in the 1970s and 1980s, and political transition in former Soviet-bloc countries at the end of the Cold War, suggests that global forces impacting multiple countries can serve as a trigger for intrastate conflict. A common conjecture is that economic forces have been a primary impetus for such episodes. In the first chapter, I analyze the effects of worldwide commodity price fluctuations in generating political conflict in developing countries. I develop a simple model to show that shocks to both the level and uncertainty of commodity export prices can elicit conflict events in developing countries. Econometric evidence from a dataset combining major intrastate political resistance campaigns and global food commodity price data lends support to this hypothesis.

In the second chapter, I examine whether corruption within one country affected by corruption within another. Understanding the interactions between political-economic

culture across countries can allow us to better grasp the implications of greater global and regional integration in recent history. Until now, few studies have examined this question in detail due to the difficulty of measuring corruption and paucity of consistent data over an adequate time span. I use a panel dataset of countries in Asia, Latin America, and the Middle East over a span of fifteen years to examine how domestic corruption reacts to the culture of corruption in the region in which the country is located. Contrary to the results of past literature, I find evidence that a reduction in regional corruption can actually lead to a worsening of corruption within a country, and vice versa. If in an open economy, regional graft lowers the level of income that a rent-seeking government can tax, a reduction in regional corruption can increase the marginal benefit of imposing a more extractive domestic policy by increasing the pool of exploitable funds. My results offer an economic reason for why corruption will be an enduring institution in a more interconnected world.

Finally, are less democratic governments more apt to intervene in the prices of imported goods than exported goods? In the third chapter of this dissertation I offer an explanation for why this might be the case, focusing on a government's choice between two alternative interventionist trade policies: import tariffs and export subsidies. If governments have incentives to exploit their political power to extract rents from citizens, they can achieve this by taxing imports rather than subsidizing exports. However, if citizens are able to discipline their governments through elections, the extent of this rent-seeking behavior can be constrained. I present a model that captures this behavior, distinguishing between the level of electoral accountability of a government and the level of bargaining power that citizens have in negotiations. Preliminary empirical evidence is presented which suggests that more authoritarian countries spend greater amounts on import tariffs than on export subsidies. These findings give insight into some of the challenges in establishing free trade amongst countries with different attitudes toward democratic institutions, both on a bilateral basis and within multilateral organizations.

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Acknowledgments

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- J.O.

Chapter 1

Price Fluctuations and Political Conflict

1.1 Introduction

Concurrent cross-country political change, such as the recent "Arab spring" revolutions in the Middle East, the experience of South American military dictatorships in the 1970s and 1980s, and political transition in former Soviet-bloc countries at the end of the Cold War, suggests that global forces impacting multiple countries can serve as a trigger for intrastate conflict. A common conjecture is that economic forces, and in particular global prices, have been a primary impetus for such episodes. For example, **The Economist** newspaper on 26 February 2011 argued that "Discontent over rising bread prices has played a part in the popular uprisings throughout the Middle East," and on 28 May 2011 noted that "Rising prices can cause mayhem....In some African markets maize and wheat prices have risen by 30% this year. Political tension invariably rises, too." In this paper, I argue that this portrayal of the effect of prices on political conflict is overly simplistic. I show that fluctuations in food commodity price levels and in commodity price uncertainty have distinct effects on civil conflict occurrence in developing countries: when a country is a net exporter of a commodity, falling (rather than rising) relative price levels lead to political tumult while an increase in uncertainty regarding the relative price of commodity exports has the same effect.

Despite the popularity of the view that movements in prices can cause political turmoil, research into this phenomenon is of limited scope, particularly in terms of measuring the interaction between price uncertainty and political constraints in developing economies. This paper seeks to fill the gap in the literature by investigating how world commodity price uncertainty and level shocks exacerbate the risk of political conflict in emerging economies. I build a small open-economy model of political competition where features of the distribution of global economic shocks can influence the domestic political process. Using a panel dataset of violent and nonviolent civil conflicts in developing countries from 1960 to 2006, I present statistical evidence in support of the model, showing that changes in both the first and second moment of the distribution of agricultural commodity export prices can predict incidence and outbreak of intrastate political uprisings. A positive change in the net export commodity price level makes conflict less likely, while a positive change in commodity relative price uncertainty works in the opposite direction. Whereas most open-economy papers focus on the welfare gains of greater integration into the global economy, by taking a first step toward modeling the interaction between domestic political constraints and global macroeconomic forces, the findings of this investigation can inform policy-makers about the extent to which a country should be insulated from international economic shocks, as well as highlight a possible channel through which global financial market reform can influence geopolitical stability.

The structure of this paper is as follows. In the next section, I discuss some of the literature related to the economics of conflict. In Section 1.3, I lay out a simple, two-period small open economy model of political conflict. Section 1.4 describes the data and presents empirical results supporting the model, and Section 1.5 concludes.

1.2 Review of Literature

Economists and political scientists have long recognized that identifying the effect of the economy on politics and vice-versa presents challenges due to the endogenous relationship between the two. The literature seeking to deal with this challenge can be broadly classified into three categories: econometric analyses of the effect of economic

variables on the likelihood of political change, econometric analyses of the effect of political events on economic variables, and theoretical modeling of economic agents' behavior under political constraints.

This paper is most closely related to the literature seeking to identify the effect of economic conditions on political processes. In an early contribution, Grier and McGarrity present empirical evidence that per-capita income growth, inflation, and unemployment have an influence on the electoral performance of incumbent politicians in the House of Representatives (Grier and McGarrity 1998). In contrast to their closed-economy approach, this paper will analyze the experience of a panel of countries in response to global economic shocks. Miguel et al. take a similar stance, and address the endogeneity of economic factors used as explanatory variables for conflict incidence (Miguel, Satyanath and Sergenti 2004). Whereas I will exploit the plausible exogeneity of world commodity prices to a small open economy, these authors use rainfall as an instrument for economic growth in African countries, finding that a five percent negative growth shock increases the likelihood of civil war by nearly one-half. Brückner and Ciccone also use commodity prices, finding that downturns in export commodity prices are associated with increased outbreak of civil war in Sub-Saharan Africa (Brückner and Ciccone 2010). In an IMF report, Arezki and Brückner find that changes in a global food price index can predict conflict in developing countries, (Arezki and Brückner 2011) but none of the papers above address the effect of uncertainty on the likelihood of civil war. A work that does address such uncertainty effects is Elbadawi and Hegre (Elbadawi and Hegre 2008). Elbadawi and Hegre examine the effect of terms of trade volatility on intrastate armed conflict, finding no robust evidence for a direct effect. One flaw of their approach may be the use of annual data; data at that level of aggregation tends to smooth away second-moment effects and understate their importance. In this work I will exploit data at monthly and quarterly frequencies in order to better capture the extent of price movements faced by developing countries.

This paper is somewhat related to the literature that seeks to econometrically identify the effect of political states on economic variables. Bittlingmayer argues that volatility in German stock prices during the decades marking the transition from Imperial to Weimar

Germany was primarily driven by the dramatic political shift (Bittlingmayer 1998). Under this assumption, politics has a clear effect on the stock market: current and past increases in volatility are associated with declines in output and stock prices. Abadie and Gardeazabal exploit the terrorist activity in Spain's Basque Country as a natural experiment to evaluate economic costs of conflict, finding that terrorist activity caused a 10 percent decrease in per-capita GDP relative to a synthetic control region and that stock prices show positive relative performance after the terrorist cease-fire (Abadie and Gardeazabal 2003). Mobarak finds a robust connection between democracy and economic development through the volatility channel: democracy and economic diversification is associated with lower volatility of macroeconomic variables, which in turn is associated with higher growth (Mobarak 2005). He uses Muslim countries to instrument for democracy to address the endogeneity of political systems with levels of economic development.

Finally, the model presented in this paper draws elements from the game-theoretic equilibrium models of interaction between economic and political agents. Marcouiller and Young show that a predatory government that increases taxation and graft at the expense of a shrinking formal economy may be acting rationally, and that the optimal level of government predation is determined by the elasticity of substitution between the goods produced in the formal and informal sectors of the economy (Marcouiller and Young 1995). Acemoglu and Robinson highlight the interaction between rich and poor agents in an unequal society and its role in generating political transitions and consolidation of democracy, finding that societies with higher asset inequality between citizens are more prone to social conflict and fiscal policy volatility (Acemoglu and Robinson 2001). Aghion et al. study the social dilemma of simultaneously desiring an effective leader while needing to limit the power of the political authority in a model where the insulation of leaders (defined as the share of votes that can block a leader's actions ex post) is determined endogenously (Aghion, Alesina and Trebbi 2004). The authors show that political insulation is positively correlated with several measures of ethnic fractionalization and polarization. In the equilibrium of Yared's dynamic model of a rent-seeking politician and citizens, temporary economic shocks can generate volatile and persistent

changes in the tax rate, demonstrating that the assumption of a benevolent government which is prevalent in the tax literature is far from benign; political distortions can create incomplete asset market-resembling behavior in an economy with complete markets (Yared 2010).

Though they may not necessarily tie their results to political consequences, labor theoretic models of trade unions and industrial strikes can also motivate studies of civil conflict. Ashenfelter and Johnson develop a model that highlights the bargaining process between trade union leaders, union members, and firm managers in determining the conditions under which a firm will incur a strike, the length of strikes, and the determination of wages (Ashenfelter and Johnson 1969). Espinosa and Rhee develop a repeated game model of wage bargaining that reconciles the Pareto inefficient outcomes of the classic monopoly union model with the efficient bargaining model (Espinosa and Rhee 1989). In their paper, strikes can emerge as part of a punishment mechanism when the firm deviates from an efficient, cooperative equilibrium. By recasting firm managers as government and unions as an organized political opposition, such models may conceivably be used to study political negotiation and failures thereof. For the purposes of this exercise however, I will abstract from the organizational aspects of and bargaining interactions between political agents.

In contrast to the general equilibrium models discussed in the political economy and labor economics literature described above, the model in this paper is a very simple, partial-equilibrium model designed to motivate the empirical exercise which follows. Details of the model are presented in Section 1.3.

1.3 Model

The model is a simplified version of the framework in (Barro 1973) featuring a two-period, small open economy populated by a representative citizen and a rent-seeking dictator. In each period, the citizen receives an exogenous, deterministic endowment m of a commodity which is subject to taxation by the dictator. After paying this tax, the citizen exports the remainder of his endowment to the rest of the world, receiving a final consumption good as payment. I assume that financial markets in the country are

underdeveloped so that the citizen is unable to access vehicles for saving or borrowing and must consume the entirety of his income after trade. The real prices of these goods (in units of the imported consumption good) are set in the rest of the world and treated as the realization of an exogenous stochastic process from the perspective of the agents in the small open economy. The dictator, in the meantime, chooses a tax rate to maximize rents in each period. If the tax rate is excessively high, the citizen is able to overthrow the dictator, at which point a new, albeit identical, dictator comes into power and the game is played again.

Within each period, the timing of events is as follows:

1. Citizens receive endowment m.
2. Citizens set reservation utility for complying with dictator. Dictator sets tax rate τ.
3. Prices are realized and goods are sold.
4. Citizens either comply with dictator policy or ignite revolution. Dictator receives rents and, if not kicked out, the value of holding office. If dictator is deposed, a new, identical authoritarian government comes into power.

1.3.1 Information

Commodity export prices in units of consumption are determined exogenously and are assumed to follow a log-normal distribution so that $\ln P_t \sim \mathcal{N}(\mu, \sigma)$, or equivalently, $P_t \sim \Lambda(\mu, \sigma)$. As the consumption good must be imported, the real commodity price P_t can also be interpreted as the small open economy's terms of trade.

1.3.2 Preferences

Prior to the realization of export prices in the first period, citizens determine their political attitudes (i.e. willingness to tolerate a government's rent-seeking behavior) in a way that maximizes their expected intertemporal utility function:

$$E_0[U_1] = E_0 \left[\ln C_1 + \beta \ \ln C_2\right] \tag{1.1}$$

subject to the budget constraint

$$C_t \leq P_t(1-\tau)m \tag{1.2}$$

where $t \in \{1, 2\}$.

At the end of each period, citizens choose whether or not to comply with the dictator based on whether their intertemporal utility is above a certain reservation level $\overline{\omega_t}$. Citizens only foment revolution to punish the dictator ex-post for failing to achieve a minimum standard of living in the country; as the dictator's possible replacement is identical from the viewpoint of the citizens, imposing this punishment is weakly optimal. As stated earlier, citizens set their reservation utility before export prices have been realized. The compliance decision is given by:

$$\pi_t = \begin{cases} 0, & \text{if } U_t < \overline{\omega_t} \\ 1, & \text{if } U_t \geq \overline{\omega_t} \end{cases} \tag{1.3}$$

The dictator's objective is to maximize rents in each period. I assume a short-sighted dictator who does not behave intertemporally, but who nevertheless seeks to hold onto office because of the presence of an exogenous value of being in power. The dictator's period utility function is:

$$E_t\left[v_D\right] = \ln r + \pi_t \ln R \tag{1.4}$$

where $\ln R$ reflects the exogenous utility of holding office. The dictator's budget constraint equates rents to the taxed portion of citizens' income:

$$r = \tau m \tag{1.5}$$

The dictator can behave in one of two ways: he may try to stay in office forever or impose maximum taxes and accept that he will be thrown out of power. In order to ensure that the dictator tries to hold onto office rather than voluntarily relinquishing power, I assume the following condition holds:

Assumption 1.

$$\ln r + \ln R \geq \ln m \quad \Rightarrow \quad r \geq m/R \tag{1.6}$$

Because the maximum amount of rent that the dictator can extract is the citizens' total endowment m, it is natural to further assume that $R > 1$, which also implies the utility of holding power is strictly positive.

1.3.3 Equilibrium

Now, given that the dictator wants to both hold onto power and maximize rents, he will set taxes as high as possible such that citizens will be expected to comply with his policies rather than ignite revolution. In other words, he sets taxes so that citizens' expected utility is arbitrarily close to their reservation level for compliance, and (1.3) holds as an equality. Combining (1.3) with (1.1) and the distribution of world prices yields the condition:

$$\overline{\omega_1} = (1+\beta)\left[\ln(1-\tau)m + \mu\right] \tag{1.7}$$

Applying the dictator budget constraint (1.5), it is possible to derive one of the conditions for the equilibrium level of rents:

$$r = m - \exp\left(\frac{\overline{\omega_1}}{1+\beta} - \mu\right) \tag{1.8}$$

Citizens will want rents to be as small as possible. To that end, they will set their reservation utility level in order to make dictator power constraint (1.6) hold with equality.

$$\ln r + \ln R = \ln m \;\Rightarrow\; r = \frac{m}{R} \tag{1.9}$$

Equations (1.8) and (1.9) will pin down the reservation utility as a function of model parameters.

At this point, prices are realized and the citizens choose whether to comply with or

revolt against the dictator. Citizens' end-of-period expected intertemporal utility is:

$$\begin{aligned} U_1 &= \ln P_1(1-\tau)m + \beta E_1 \left[\ln P_2(1-\tau)m\right] \\ &= \ln P_1(1-\tau)m + \beta \left[\ln (1-\tau)m + \mu\right] \\ &= \ln P_1(1-\tau)m + \beta \left[\ln (1-\tau)m + \ln E_1[P_2] - \tfrac{\sigma^2}{2}\right] \end{aligned} \tag{1.10}$$

The last two lines in (1.10) follow from the properties of the log-normal distribution. Suppose there is a permanent, positive shock to the mean of the log relative price distribution at the end of period 1, which could happen in the realistic case that the log of prices is non-stationary. By the second line of (1.10), the citizens' end-of-period expected utility is increased, making it more likely the citizens will achieve their reservation utility level and, consequently, comply with the dictator. Suppose in period 1, there is a positive, mean-preserving, permanent shock to the variance of the log relative price distribution (i.e. P is held constant). By the third line of (1.10), this shock to uncertainty lowers utility and makes it more likely that revolution will occur. This analysis suggests that both first-moment and second-moment relative price shocks should play a role in determining conflict in developing countries.

The second period of the model is solved in similar fashion. Realized utility at the end of the terminal period can be expressed as: $U_2 = \ln P_2 + \ln (1-\tau)m$. However, as there are no expectations of future utility at the end of the period, ultimately only negative shocks to the relative price level can send a country into conflict.

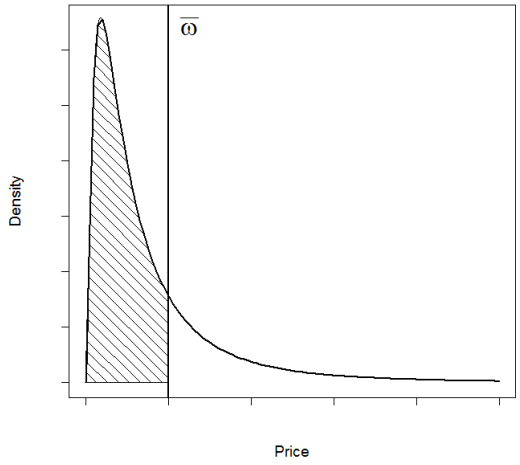
1.3.4 Comparative Statics

Using (1.8) and (1.9), the solution for the reservation level in period 1 is:

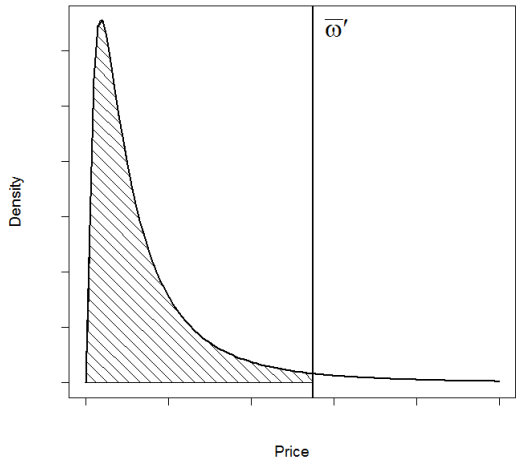
$$\overline{\omega_1} = (1+\beta)\left\{\ln m + \ln\left(\frac{R-1}{R}\right) - \mu\right\} \tag{1.11}$$

Taking the derivative of (1.11) with respect to the level of endowment, m, yields: $\dfrac{\partial \overline{\omega_1}}{\partial m} = \dfrac{1+\beta}{m} > 0$. As citizens' income increases (for a small enough increase such that the dictator power-holding constraint (1.6) holds), they demand a higher level of utility in order to comply with the dictator's policies. For a given distribution of prices, this

implies that richer countries are more likely to overthrow a dictator. Figure 1.1 depicts the reservation utility level calculated for low and high values of m, superimposed on the probability density curve for the relative price. A realization of relative price levels lower than the expected value used in the calculation of the reservation utility can result in political conflict. The shaded areas under the curves in Figure 1.1 represent the probability of revolution. An increase in m increases the size of the region in which prices are associated with conflict and therefore increases the likelihood of revolution.



(a) Lower m



(b) Higher m

Figure 1.1: Effect of increase in m

Similarly, taking the derivative of (1.11) with respect to the exogenous value of

holding power, R, yields: $\frac{\partial \overline{\omega_1}}{\partial R} = \frac{1+\beta}{R(R-1)}$. This derivative is positive so long as $R > 1$, or equivalently, so long as the dictator receives positive utility from being in office. As the value of holding office increases, the dictator is able to request a lower tax rate from the citizens; combining equilibrium conditions (1.5) and (1.9) shows that the equilibrium tax rate is $\tau^* = \frac{1}{R}$. However, citizens do not see this lower tax rate as a sign of the dictator's beneficence. Rather, it arises in equilibrium because the citizens, who are aware of the higher value of holding office, take the liberty to impose even more stringent demands upon their government. This demand takes the form of a higher reservation level of utility for not revolting, and by the same logic as before, increases the likelihood that the dictator will be overthrown, for a given distribution of prices.

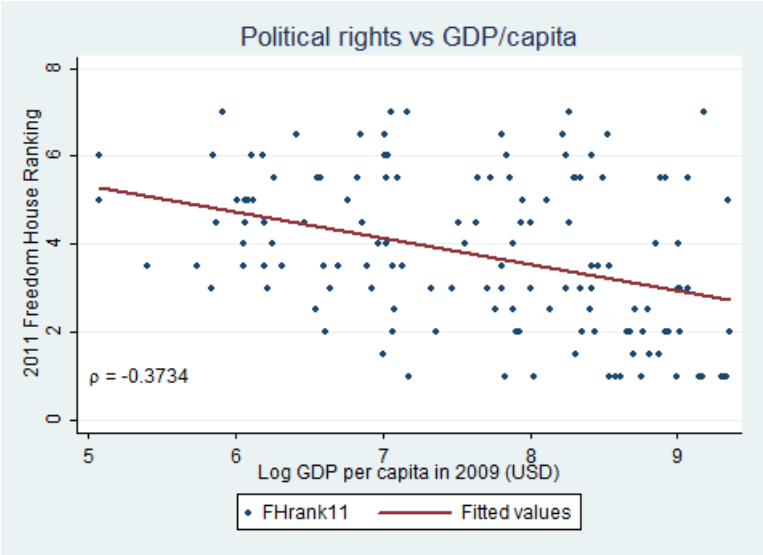
These results find some support in the data. In the graphs below, I use the 2011 Freedom in the World survey's combined average rating of political rights and civil liberties[1] to proxy for whether or not a country has overthrown a dictator. Rating is done on a scale from 1 to 7, with lower scores indicating higher measures of political rights and civil liberties. Presumably, countries which have authoritarian governments should garner a higher ranking. 2009 per-capita GDP data taken from the World Bank's *World Development Indicators* is used as a measure of income. In Figure (1.2a), I drop all countries which are classified as high-income by the World Bank in 2009.[2] The model suggests that we should observe a negative relationship between level of freedom and income, a prediction which is borne out in the data.

Measuring the exogenous value of being in power proves to be challenging. One reason a dictator might choose to hold power is to enjoy the geopolitical prestige offered by international recognition and influence. I conjecture that international influence can partly be explained by military power, especially to the extent that military clout can be projected beyond domestic borders, and thus sophistication of a nation's armed forces may serve as a reasonable measure of the external utility of holding office. I use the ratio of military expenditure to military personnel in 2007, taken from the Correlates of War Project's *National Material Capabilities v4.0* dataset, to measure military sophistication.

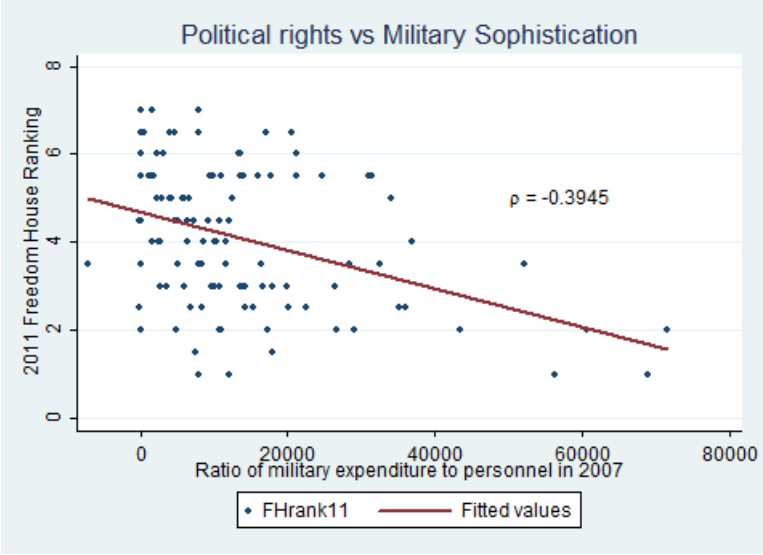
[1] Published by Freedom House (see appendix).

[2] Technically, the World Bank definition categorizes countries according to threshold limits for gross national income. Here I apply the same thresholds to gross domestic product.

In Figure (1.2b), I drop all high-income countries in the year 2007, as categorized by the World Bank. The model indicates that we should expect a negative relationship between military sophistication and the measure of political freedoms, which proves to be the case.



(a) Freedom index vs GDP



(b) Freedom index vs military expenditure

Figure 1.2: Relationship between political system & measures of income and value of office-holding

1.4 Empirical Results

1.4.1 Data

The primary implication of the model in the previous section is that both level and uncertainty shocks to commodity export price levels can generate conflict situations within countries. To test this proposition, I construct a panel dataset of all countries between 1960 and 2006, at the quarterly frequency. For each quarter-country pair, I record whether or not the country has been in internal conflict as recorded by the Nonviolent and Violent Conflict Outcomes (NAVCO) dataset compiled by Stephan & Chenoweth (Stephan and Chenoweth 2008). This unique dataset takes into account not only violent, armed resistance campaigns, but non-violent revolutions as well, and considers only major events (>1,000 battle deaths in violent conflicts or >1,000 participants in nonviolent conflicts) with maximal objectives (posing a major challenge to the existing order, producing a major government crackdown on participants or resulting in the expulsion of a foreign occupier or domestic regime, self-determination, or secession). The NAVCO dataset is recorded at annual frequency, so I cross-reference the conflict dates with those listed in Gleditsch, et al. to narrow down the precise quarters of conflict where overlap between the databases occurs (Gleditsch, Wallensteen, Eriksson, Sollenberg and Strand 2002). When no overlap occurs, I assume the country has been experiencing conflict for the entire year. This is a reasonable assumption due to the difficulty of identifying precise dates of conflict onset and resolution; in some cases the two databases disagree on even the year that a conflict started or ended, indicating the blurred distinction between whether an incident is an isolated event or part of a major campaign. Formally, the dependent variable records conflict incidence as:

$$\text{conflict}_t = \begin{cases} 1, & \text{if conflict in period t} \\ 0, & \text{if no conflict in period t} \end{cases} \tag{1.12}$$

Descriptive statistics for the conflict variable are given in Table 1.1.

To measure prices, I use international commodity price data from the *World Bank GEM Commodities* database. (Bank 2011) The data used are given at monthly frequency

Variable	Obs	Mean	Std. Dev.	Min	Max
conflict	21808	0.177	0.382	0.000	1.000

Table 1.1: Summary statistics for conflict variable, quarterly frequency

from January 1960 to December 2006. I use the constant price (in year 2000 USD) series for 21 agricultural commodities, normalizing the initial monthly observation to unity. Quarterly means were used to measure the commodity's price level. For price volatility I calculate the quarterly mean of monthly standard deviations for each commodity.[3] I then construct aggregate commodity price level (PI) and volatility (VarPI or Vol_xPI) indices for each country by taking a weighted sum of commodity prices or volatilities respectively, where the weights are a country's time-invariant net export shares of each commodity. If a country is not a net exporter of a particular commodity, that commodity receives a weight of zero in the country's export price and volatility indices. For export shares I use data on commodity imports and exports in 2006 from the United Nations Commodity Trade Statistics Database.

A list of commodities and accompanying summary statistics at the quarterly frequency appears in Table 1.2, along with summary statistics for the aggregate commodity export price level and uncertainty indices. Vol_1PI and Vol_2PI are alternative uncertainty series whose construction will be described in the robustness checks section of this paper. Indices are equal to zero in countries that are not net exporters of any of the 21 commodities; in these cases the countries are dropped from the sample due to perfect collinearity between the constant term and the constant commodity price level and uncertainty indices. In the estimated equations that follow, level and uncertainty indices are differenced and filtered of time effects by regressing them on a set of quarterly time dummies.

Finally, I restrict the sample of countries in the estimation by partitioning them into high-, middle-, and low-income countries according to the 1987 World Bank *World Development Indicators* definition. High income and upper middle income countries are dropped from the sample, as these countries are most likely to violate the small

[3]Monthly standard deviations are calculated as a rolling standard deviation of commodity prices where the moving window includes the first to third lag. Results are similar when calculated for as many as six lags of monthly price observations.

Variable	Obs	Mean	Std. Dev.	Min	Max
banana	21808	0.826	0.170	0.481	1.382
barley	21808	1.058	0.214	0.532	1.872
cocoa	21808	0.830	0.434	0.332	2.681
coconut oil	21808	0.524	0.259	0.162	1.572
coffee	21808	0.885	0.519	0.184	3.556
copra	21808	0.510	0.246	0.167	1.544
groundnut oil	21808	0.768	0.252	0.364	1.679
maize	21808	0.806	0.287	0.388	1.661
beef	21808	0.980	0.301	0.487	1.956
chicken	21808	0.984	0.149	0.674	1.278
orange	21808	1.024	0.252	0.604	2.006
palm oil	21808	0.669	0.282	0.249	1.671
rice	21808	0.904	0.444	0.388	2.877
sorghum	21808	0.877	0.297	0.453	1.893
soybean meal	21808	0.753	0.301	0.355	2.540
soybean oil	21808	0.838	0.353	0.370	2.393
soybean	21808	0.875	0.306	0.485	2.391
sugar	21808	1.112	0.386	0.796	4.566
tea	21808	0.540	0.199	0.270	1.000
tobacco	21808	0.497	0.162	0.288	1.000
wheat	21808	0.771	0.260	0.414	1.944
PI	14100	0.821	0.340	0.162	4.566
VarPI	14025	0.062	0.095	0.000	1.255
Vol_1PI	14025	0.017	0.068	0.000	1.576
Vol_2PI	14100	0.065	0.093	0.000	1.773

Table 1.2: Summary statistics for prices and uncertainty series, quarterly frequency

open economy assumption and for whom domestic supply and demand effects are likely to influence global commodity prices. For the remaining 75 countries in the sample I estimate the effect of an increase in the change of the price index and the uncertainty index on the probability of conflict incidence. A preview of the primary result can be seen in the cross-sectional illustration of Figure 1.3. In this figure, the average conflict incidence is plotted against the average change in the price uncertainty index over the 1960s, with each point representing a country. The correlation coefficient is positive and significant at 0.2512, suggesting there may be some connection between higher price uncertainty and conflict incidence. Results for the estimation of the dynamic models over the full sample period are shown beginning with Table 1.3.

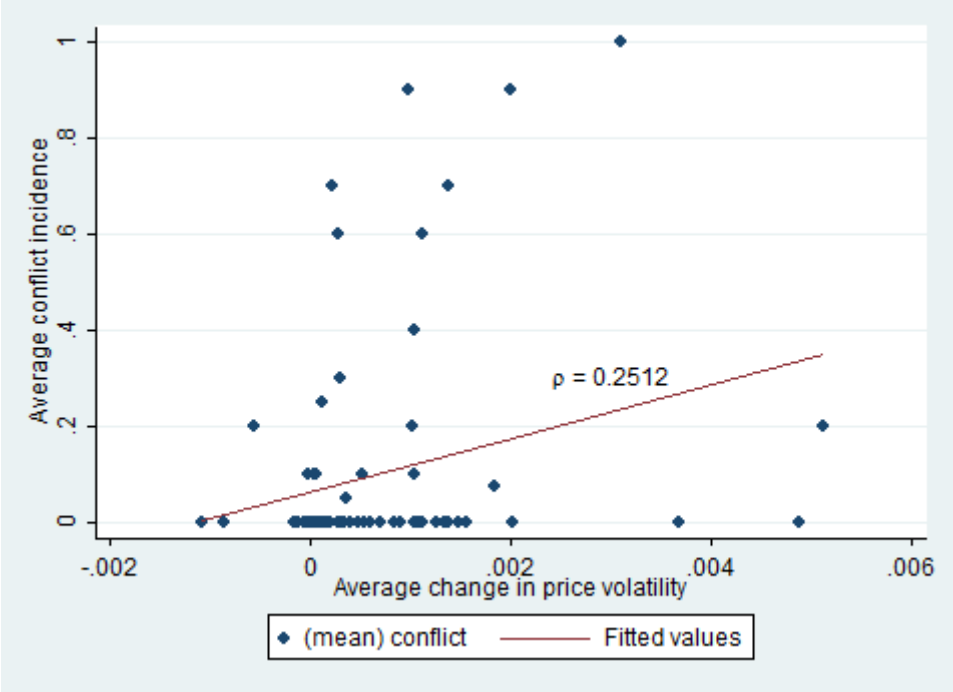


Figure 1.3: Average conflict incidence versus average change in uncertainty, 1960s

1.4.2 Methodology and Estimation

In the following tables I present linear and nonlinear probability models which estimate the effect of changes in price level and uncertainty on the likelihood of conflict incidence. Incorporating changes (rather than levels) of the commodity price level and uncertainty indices reflects the fact that the realizations of prices and uncertainty only matter for

conflict insofar as they lower utility relative to the reservation level of utility associated with keeping the dictator in power. The reservation utility level is, in turn, calculated from the expected value of prices and uncertainty conditional on the information set of the previous period. Though the model in Section 1.3 featured a stationary distribution for the mean and standard deviation of prices, it is more empirically likely that commodity price levels and uncertainty are nonstationary.[4] For example,

$$\ln P_t = \ln P_{t-1} + \sigma_{t-1}\eta_t \tag{1.13}$$

$$\sigma_t = \sigma_{t-1} + \epsilon_t \tag{1.14}$$

where $\eta_t \sim \mathcal{N}(0,1)$, ϵ_t is a white-noise shock, and $\{P_0,\ \sigma_0\}$ are given. In this case, citizens' expected utility prior to the realization of commodity price level and uncertainty shocks can be written as:

$$E_0[U_1] = \overline{\omega_1} = \ln P_0 + \beta \ln E_0[P_1] - \frac{\beta\sigma_0^2}{2} + (1+\beta)\ln\{(1+\tau)m\} \tag{1.15}$$

Utility after the realization of prices in period 1 can be written as:

$$U_1 = \ln P_1 + \beta \ln E_1[P_2] - \frac{\beta\sigma_1^2}{2} + (1+\beta)\ln\{(1+\tau)m\} \tag{1.16}$$

Conflict occurs when the difference between (1.16) and (1.15) is less than zero:

$$\text{conflict}_1 = \begin{cases} 1, & \text{if } \Delta \ln P_1 + \beta\left(\ln E_1[P_2] - \ln E_0[P_1]\right) - \frac{\beta}{2}\Delta\sigma_1^2 < 0 \\ 0, & \text{otherwise} \end{cases} \tag{1.17}$$

Equation (1.17) illustrates that, *ceteris paribus*, negative changes in the real export price level and positive, mean-preserving changes in price uncertainty are likely to result in conflict. Equation (1.17) also demonstrates that the predictions of the theoretical model can be readily assessed using a binary outcome model such as the logit or probit model,

[4] Heteroskedasticity-robust Hadri Lagrange multiplier tests for the stationarity of the net export commodity price index and the three uncertainty indices used in this paper strongly reject the null hypothesis that all panels are stationary ($p < 0.0001$), regardless of inclusion of time trends or removal of cross-sectional means from the respective series.

including changes in price level and uncertainty as regressors.

Results

	LPM-RE	LPM-FE	LOGIT-RE	LOGIT-FE	PROBIT-RE
ΔPI_t	-0.028+	-0.028+	-0.350+	-0.350+	-0.203+
	(0.017)	(0.017)	(0.198)	(0.198)	(0.114)
ΔVarPI_t	0.082**	0.082**	0.952**	0.952**	0.556**
	(0.029)	(0.029)	(0.321)	(0.321)	(0.194)
Constant	0.194***	0.194***	-2.999***		-1.585***
	(0.029)	(0.000)	(0.422)		(0.214)
$\ln \sigma_u^2$			2.221***		0.833***
			(0.232)		(0.229)
N	13950	13950	13950	10230	13950
Clusters	75	75	75	55	75
Overall p	0.007	0.010	0.005	0.006	0.005
Hausman p	0.481				

$+ \ p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.3: Effect of price level and volatility on conflict incidence

Table 1.3 shows the effect of a change in the price and volatility indices on the likelihood of conflict for the baseline linear probability model (with both fixed and random effects for countries), the logit model with both fixed and random effects, and the random-effects probit model.[5] Across all specifications of the model, I find that an increase in the export price level is negatively associated with the incidence of conflict, as predicted by the theoretical model. The price level effect is significant at the 10% level in all specifications of the model. Strikingly, even after controlling for price levels I not only find that price volatility remains significant, but that it appears to be a more powerful predictor of conflict incidence than the price level effect. The price uncertainty effect is significant at $\alpha = 0.01$ in all the estimated models, and its coefficient is of the correct sign. An increase in price volatility over the previous quarter is associated with an increased probability of internal conflict. These findings lend support to the conclusions of the model and stand in contrast to works which downplay the effect of

[5] As a Hausman test for fixed- versus random-effects does not definitively reject the assumptions of the random effects model, I report both specifications. $P = 0.4813$.

price uncertainty on political stability.

While examining the response of conflict to changes in export price levels and uncertainty in isolation may be indicative of the true direction of the effects of these variables, the model in Section 1.3 suggests that the relevant level and uncertainty variables pertain to export prices relative to the price of the imported consumption good. This suggests that the estimation in Table 1.3 should incorporate a measure of consumption or welfare. For example, if a welfare-deteriorating decrease in export prices is offset by a welfare-augmenting decrease in the prices of other goods in the consumption basket, there can zero net effect on citizens' utility and the likelihood of conflict incidence. If the prices of the non-exported consumption goods fall enough relative to falling export prices, so that there is a net increase in utility, the conclusions of the model may even be reversed when looking at the data. Although the presence of such effects would actually bias the econometric results against the predictions of the model, the estimation presented in Table 1.3 can be more tightly linked to the interpretation of export prices as relative prices by controlling for changes in the standard of living within a country. Additionally, both the model in Section 1.3 and the previous literature on the economics of civil conflict allude to an important role for income in determining conflict incidence. Formally including income fluctuations in the estimation will show that the results in Table 1.3 are robust to incorporating this additional explanatory variable. In the tables that follow, I augment the baseline regression with several different measures of income.

As reliable quarterly national accounts data are incomplete or unavailable for many of the countries in my sample, I proxy for income with rainfall in the spirit of Miguel et al.'s use of rainfall as an instrument for income shocks (Miguel et al. 2004). Data on rainfall are from the Global Historical Climatology Network version 2-Monthly dataset. Quarterly rainfall series are expressed in meters and are obtained for each country by averaging rainfall levels across weather stations in each country and summing by quarter. The results in Table 1.4 are largely similar to those in Table 1.3. When changes in rainfall are considered as proxies for income changes, the coefficient on price volatility remains positive and significant at the 1% level. The coefficient on price level changes remains negative significant at the 10% level in all specifications. Rainfall itself does not appear

to affect the likelihood of conflict incidence.

	LPM-RE	LPM-FE	LOGIT-RE	LOGIT-FE	PROBIT-RE
ΔPI_t	-0.034+	-0.034+	-0.403+	-0.403+	-0.229+
	(0.018)	(0.018)	(0.206)	(0.206)	(0.119)
ΔVarPI_t	0.088**	0.088**	0.996**	0.996**	0.570**
	(0.032)	(0.032)	(0.336)	(0.336)	(0.205)
Δrain_t	0.002	0.002	0.018	0.018	0.012
	(0.001)	(0.001)	(0.015)	(0.015)	(0.009)
Constant	0.198***	0.198***	-2.873***		-1.528***
	(0.031)	(0.000)	(0.407)		(0.207)
$\ln \sigma_u^2$			2.207***		0.838***
			(0.240)		(0.238)
N	12621	12621	12621	9580	12621
Clusters	74	74	74	55	74
Overall p	0.014	0.034	0.007	0.008	0.007
Hausman p	0.259				

$+ \ p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.4: Effect of price level, volatility, and income proxy (rainfall) on conflict incidence

The previous findings continue to hold when international reserves augment the regression as a proxy for income. Data on international reserves are from the IMF's *International Financial Statistics* and are measured in billions of SDR. Because governments might adjust international reserve holdings in order to pay for the expense of waging conflict, in the first column of Table 1.5, I instrument the current change in international reserves with the current and two lags of changes in rainfall, following the literature's use of weather conditions as exogenous instruments for economic variables. The IV specification shows that a positive change in the income proxy is associated with a statistically insignificant decrease in the probability of conflict incidence and that after controlling for the income proxy, changes in price level and volatility remain significant at the 5% level. The price effects are in the direction predicted by the model. Table 1.5 reports that rainfall is a valid instrument for reserves in the test of overidentifying restrictions.[6] However, a test of the hypothesis that international reserves can be considered an exogenous regressor is unable to reject its null hypothesis. Accordingly, I

[6] A test of overidentifying restrictions in the linear probability model is equivalent to the Hausman test for fixed versus random effects. Here I am unable to reject the random effects model.

report OLS, logit and probit estimates of the effect of price levels and uncertainty on probability of conflict while controlling for changes in international reserves. Higher reserves are associated with lower conflict incidence in all these models, but are significant at the 10% level in the linear probability models only. The results for price fluctuations are similar across all specifications and conform to the theoretical model in Section 1.3. When international reserves holdings proxy for income, exogenous changes in commodity export prices and volatility are still significantly correlated with conflict incidence in lower income countries. Higher export prices make conflict less likely, while higher volatility makes conflict more likely.

	IV-LPM-FE	LPM-RE	LPM-FE	LOGIT-FE	PROBIT-RE
Δreserves$_t$	-0.033	-0.005+	-0.005+	-0.057	-0.033
	(0.038)	(0.003)	(0.002)	(0.201)	(0.112)
ΔPI$_t$	-0.048*	-0.046*	-0.046*	-0.563**	-0.328**
	(0.020)	(0.019)	(0.019)	(0.201)	(0.117)
ΔVarPI$_t$	0.089*	0.078*	0.078*	0.915*	0.531*
	(0.040)	(0.036)	(0.036)	(0.386)	(0.233)
Constant		0.180***	0.209***		-1.882***
		(0.031)	(0.000)		(0.289)
$\ln \sigma_u^2$					1.166***
					(0.310)
N	9399	10429	10429	7222	10429
Clusters	69	73	73	45	73
Overall p	0.028	0.003	0.013	0.007	0.005
Over-i.d. test p	0.267	0.281			
Endog. test p	0.317				

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.5: Effect of price level, volatility, and income proxy (international reserves) on conflict incidence

An alternative to finding proxies for quarterly income data in less developed countries is to impute higher frequency data from more the complete data series available at the annual frequency. This is what I do in Table 1.6. I construct annual income per capita series using data on household consumption expenditure and population from the UN. I then interpolate this series to the quarterly frequency using the proportional Denton method described by the IMF (Bloem, Dippelsman and Maehle 2001). This

technique constructs a time series of quarterly national accounts estimates from annual observations by benchmarking the quarterly series as proportional to a higher frequency indicator via least-squares minimization. For this application, quarterly GDP in the OECD aggregate is used as the higher-frequency indicator for the interpolated data. Under the assumption that variation in total GDP of the OECD is caused by international macroeconomic forces that affect middle- and lower-income countries in similar fashion, the interpolated income per capita series will be highly correlated with actual quarterly income per capita. The Denton method imposes the constraint that the interpolated series must aggregate to the annual frequency totals, minimizing the risk that inaccurately interpolated values contaminate the series to the extent that it no longer reflects actual income per capita.

As a rough check of the accuracy of the Dentonized series, I take existing quarterly GDP series from the IMF *International Financial Statistics*, recorded in year 2000 constant prices (in national currency units), and divide by the year 2000 US dollar exchange rate (specified national currency units / USD). I divide by quarterly population to compute a quarterly GDP per capita series in U.S. dollars. Quarterly population was obtained by linearly interpolating annual population data from the UN. The coefficient of correlation between the IMF series and the Dentonized measure of income per capita was significant and positive ($\rho = 0.7010$), suggesting reasonable accuracy when using the interpolated series. However, the IMF series had a scant 152 observations, as opposed to 14,932 when using the interpolated measure.

Table 1.6 indicates that after controlling for the interpolated measure of income, the uncertainty effect of commodity export prices on conflict incidence is still present. The first column instruments for percentage change in income per capita with four lags of rainfall shocks, but as I am unable to reject the hypothesis that the income measure is exogenous, I also report OLS, logit, and probit results. Though price level and income changes changes have the expected sign in all columns of Table 1.6, the estimated coefficients are not significant. However, all the columns indicate that higher price volatility is correlated with a higher probability of conflict incidence, an effect that is significant at the 5% level.

	IV-LPM-FE	LPM-RE	LOGIT-RE	LOGIT-FE	PROBIT-RE
ΔPI_t	-0.026	-0.022	-0.278	-0.278	-0.157
	(0.019)	(0.018)	(0.215)	(0.214)	(0.125)
ΔVarPI_t	0.081*	0.084*	0.976*	0.974*	0.591**
	(0.036)	(0.033)	(0.384)	(0.382)	(0.223)
Δincome_t	-0.249	-0.299	-2.425	-2.383	-1.309
	(3.288)	(0.208)	(1.939)	(1.905)	(0.964)
Constant		0.227***	-2.965***		-1.645***
		(0.034)	(0.585)		(0.319)
$\ln \sigma_u^2$			2.777***		1.416***
			(0.343)		(0.343)
N	9131	10385	10385	6576	10385
Clusters	73	75	75	48	75
Overall p	0.125	0.037	0.061	0.066	0.044
Over-i.d. test p	0.185	0.316			
Endog. test p	0.490				

$+ \ p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.6: Effect of price level, volatility, and income on conflict incidence

1.4.3 Robustness Checks

Alternative Measures of Volatility

In Tables 1.7 & 1.8, I show that the econometric results obtained in the previous tables are somewhat robust to alternative measures of price uncertainty. For each commodity-month pair, I generate a measure of price volatility by taking a rolling variance of monthly commodity prices, where the moving window is comprised of the first to the third lag of the commodity price. This yields a monthly commodity price volatility index which is then converted to the quarterly frequency by taking means. I then generate a country-specific price uncertainty index (Vol_1PI) by aggregating these commodity-level volatility indices with the same net export weights used in the baseline results. Table 1.7 presents least-squares estimates demonstrating that changes in this index remain positive and significant at the 10% level after controlling for changes in price level and rainfall shocks. Furthermore, the coefficient on export price level changes is negative and significantly associated with probability of conflict incidence at the 5% level. In the IV specification controlling for changes in international reserves (where instruments

were current and lagged changes in rainfall levels), an increase in the change in price level lowers probability of conflict at the 1% level of significance, while an increase in the change in price volatility is associated with a higher probability of conflict at the 10% level. When percentage change in income per capita is included in the regression, the estimated effects for price level and uncertainty changes are similar. In the fourth column of Table 1.7, the instrument set was comprised of four lags of changes in rainfall levels; the coefficient on volatility is positive and significant at the 5% level while the price level coefficient is negative and signficant at 10%. I also report fixed-effect logit and random-effects probit specifications where income per capita is treated as exogenous. The results show that all price effects are significant at the 10% level and are in the direction predicted by the model.

	LPM-RE	LPM-RE	IV-LPM-FE	IV-LPM-FE	LOGIT-FE	PROBIT-RE
ΔPI_t	-0.036*	-0.043*	-0.060**	-0.041+	-0.404+	-0.217+
	(0.018)	(0.019)	(0.021)	(0.022)	(0.221)	(0.128)
$\Delta \text{Vol}_1\text{PI}_t$	0.148+	0.160+	0.177+	0.186*	1.980+	1.035+
	(0.085)	(0.089)	(0.107)	(0.092)	(0.994)	(0.560)
Δrain_t		0.002				
		(0.001)				
$\Delta \text{reserves}_t$			-0.029			
			(0.037)			
Δincome_t				-0.194	-2.380	-1.308
				(3.328)	(1.902)	(0.963)
Constant	0.194***	0.198***				-1.645***
	(0.029)	(0.031)				(0.319)
$\ln \sigma_u^2$						1.416***
						(0.343)
N	13950	12621	9399	9131	6576	10385
Clusters	75	74	69	73	48	75
Overall p	0.077	0.074	0.039	0.133	0.086	0.096
Over-i.d. test p	0.421	0.481	0.258	0.224		
Endog. test p			0.367	0.449		

$+\ p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.7: Estimation using price variance to measure price uncertainty

Table 1.8 presents evidence that export price uncertainty increases the probability of civil conflict occurring when the quarterly range of prices, rather than standard deviation or variance, is used as a measure of export price uncertainty. For each commodity-quarter pair, I generate a measure of volatility by taking the range of the monthly commodity prices in that quarter. The price volatility index Vol_2PI is generated by

taking a weighted sum of these quarterly range series, using the same net export weights described in the previous section. This index remains significant at the 5% level in the baseline model as well as the model that controls for changes in rainfall. In the third column of Table 1.8, changes in international reserves are instrumented by two lags of changes in rainfall. Price volatility is positive and significant at the 1% level when the instrumental variables approach is used, as well as in the logit and probit specifications which treat international reserves as exogenous. In all columns, the price level effect is in the direction predicted by the model, but is only significant when changes in reserves are included in the regression. In the IV specification of the model that controls for changes in income per capita (instrumented by current and two lags of changes in rainfall level), there is evidence of a positive association between price volatility and conflict incidence at $\alpha = 0.10$. Neither rainfall, reserves, nor income per capita appear to have a significant effect on the probability of conflict incidence when price volatility is expressed as a quarterly price range.

	LOGIT-RE	LOGIT-RE	IV-LPM-FE	LOGIT-FE	PROBIT-RE	IV-LPM-FE
ΔPI_t	-0.287	-0.352	-0.050*	-0.524*	-0.301*	-0.021
	(0.219)	(0.228)	(0.023)	(0.216)	(0.124)	(0.024)
$\Delta \text{Vol}_2\text{PI}_t$	0.217*	0.280*	0.040**	0.308**	0.172**	0.024+
	(0.103)	(0.128)	(0.015)	(0.104)	(0.061)	(0.015)
Δrain_t		0.015				
		(0.015)				
$\Delta \text{reserves}_t$			-0.033	-0.057	-0.033	
			(0.040)	(0.198)	(0.110)	
Δincome_t						2.152
						(3.031)
Constant	-2.980***	-2.857***			-1.873***	
	(0.418)	(0.401)			(0.284)	
$\ln \sigma_u^2$	2.203***	2.178***			1.155***	
	(0.234)	(0.239)			(0.308)	
N	14025	12694	9428	7245	10458	9126
Clusters	75	74	69	45	73	73
Overall p	0.117	0.172	0.041	0.017	0.014	0.380
Over-i.d. test p			0.494			0.392
Endog. test p			0.474			0.286

$+ \ p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.8: Estimation using price range to measure price uncertainty

Effect on Conflict Outbreak

When using the quarterly range to measure price uncertainty, there is evidence that volatility significantly impacts the outbreak of conflict as well as its incidence. To measure outbreak of intrastate conflict, I define the civil conflict onset indicator variable in quarter t as 1, if there is a conflict in t but no conflict in $t-1$; 0, if there is no conflict in quarter t; and undefined otherwise. The intuition behind the construction of this variable is as follows: if there is no conflict within a country at time t, neither is there outbreak of a conflict; if a country transitions from a peaceful state to a conflict state, conflict outbreak has occurred; transitions from conflict to conflict are recorded as undefined because it is not clear whether conflict in the latter period is a continuation of the conflict in the previous period or the outbreak of a new (though possibly related) conflict event. Although this new dependent variable causes some loss of observations due to missing values, it helps to account for possible path-dependency of likelihood of conflict. Countries which have historically undergone long episodes of civil conflict may be more likely to be in conflict in the current quarter. The previous analysis of conflict incidence only takes into account whether conflict is occurring in a particular period, treating each period as the same. In contrast, the civil conflict outbreak variable takes history into consideration by recording the emergence of new conflicts as defined by transitions from a peaceful state to a conflict state, while remaining agnostic on consecutive periods of conflict incidence.

In Table 1.9, I present the results of the instrumental variables and nonlinear models using the onset variable; results for OLS linear probability models are qualitatively similar and statistically significant. I control for income changes by including a measure of international reserves as an income proxy in the first three columns, and the proportionally interpolated measure of income per capita in the fourth through sixth columns of Table 1.9. In the first column, change in international reserves is instrumented by current and lagged change in rainfall. In this specification, as well as in the logit and probit models that include changes in reserves, an increase in the export price level over the previous period causes a decrease in the likelihood of conflict outbreak while an increase in the change in export price volatility causes an increase in the likelihood of

conflict outbreak. The fourth column of Table 1.9 instruments for change in per capita income with the current and three lags of changes in rainfall. The estimated coefficients on the price level variable and the volatility variable are of the sign predicted by the model; the level effect is significant at the 5% level and the uncertainty effect is significant at the 10% level. As I am unable to reject exogeneity of per capita income, I also report fixed-effects logit and random-effects probit specifications controlling for income changes. In the fifth and sixth column, the coefficient on price changes is negative and significant at $\alpha = 0.01$ while the coefficient on the change in volatility is positive and significant at $\alpha = 0.10$. Changes in income are also significant and of the expected sign in these two columns, lending support to the predictions of the model.

	IV-LPM-FE	LOGIT-FE	PROBIT-RE	IV-LPM-FE	LOGIT-FE	PROBIT-RE
ΔPI_t	-0.032**	-4.492***	-1.433***	-0.032*	-3.539***	-1.175**
	(0.012)	(1.177)	(0.393)	(0.013)	(0.979)	(0.345)
$\Delta \text{Vol}_2\text{PI}_t$	0.024+	4.104*	1.516*	0.025+	3.438+	1.250+
	(0.013)	(1.717)	(0.728)	(0.015)	(1.782)	(0.736)
$\Delta \text{reserves}_t$	0.027	-0.122+	-0.029			
	(0.140)	(0.070)	(0.065)			
Δincome_t				-0.960	-6.901+	-3.713+
				(1.351)	(4.010)	(1.953)
Constant			-2.493***			-2.465***
			(0.057)			(0.056)
$\ln \sigma_u^2$			-3.615			-4.205
			(4.636)			(9.082)
N	7434	4958	8340	6986	4594	8062
Clusters	67	41	72	67	44	71
Overall p	0.066	0.000	0.005	0.105	0.002	0.001
Over-i.d. test p	0.130			0.397		
Endog. test p	0.474			0.598		

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.9: Effect of price uncertainty (range) on onset of conflict

Increased volatility also appears to increase the probability of civil conflict outbreak when the standard deviation and variance are used to measure export price uncertainty. However, the evidence is sensitive to the inclusion of the income measure: I was not able to find evidence of a volatility effect when rainfall or reserves proxied for income, as these income proxies were not signficantly related to probability of conflict outbreak and inflated the standard errors of the price variables when included. Results for the IV, logit, and probit specifications are presented in Table 1.10. Linear probability models

estimated with OLS yield similar results. All estimates indicate that higher export prices are associated with significantly decreased likelihood of civil conflict outbreak, supporting the model hypothesis. Increases in per capita income are also correlated with less likelihood of conflict onset, and are significant at the 10% level in the logit and probit specifications. The first three columns report estimated coefficients when the standard deviation is used to measure price volatility. In the first column, higher volatility is associated with a higher likelihood of civil conflict outbreak, an effect that is significant at the 10% level. The next two columns treat percentage changes in per capita income as exogenous and find a volatility effect that is positive and significant at the 5% level. In the fourth column, changes in per capita income are instrumented with current and three lags of changes in rainfall level (the same instruments are used in the first column). The estimated coefficient on volatility is positive and significant at the 5% level. In the fixed-effect logit and random-effect probit specifications that control for per capita income, changes in volatility continue to have a positive and significant effect on conflict outbreak probability.

	IV-LPM-FE	LOGIT-FE	PROBIT-RE	IV-LPM-FE	LOGIT-FE	PROBIT-RE
ΔPI_t	-0.031**	-3.011**	-0.977**	-0.031**	-2.994**	-0.941**
	(0.012)	(0.943)	(0.294)	(0.012)	(0.921)	(0.299)
ΔVarPI_t	0.057+	6.122*	2.220*			
	(0.030)	(2.903)	(1.113)			
$\Delta \text{Vol}_1\text{PI}_t$				0.057*	6.128*	1.881*
				(0.028)	(2.465)	(0.938)
Δincome_t	-0.973	-6.544+	-3.688+	-0.962	-6.219+	-3.616+
	(1.352)	(3.724)	(1.904)	(1.349)	(3.567)	(1.870)
Constant			-2.463***			-2.459***
			(0.056)			(0.056)
$\ln \sigma_u^2$			-4.030			-4.075
			(9.141)			(9.151)
N	6986	4594	8062	6986	4594	8062
Clusters	67	44	71	67	44	71
Overall p	0.067	0.006	0.001	0.078	0.002	0.001
Over-i.d. test p	0.422			0.404		
Endog. test p	0.579			0.599		

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.10: Effect of price uncertainty (variance and s.d.) on onset of conflict

In a sense, the conflict outbreak indicator onset_t lies in the middle of a spectrum whose extremes depict opposing views of consecutive periods of civil strife. If instead

of recoding conflict outbreak at time t as a missing value when $\text{conflict}_t = 1$ and $\text{conflict}_{t-1} = 1$, we had recoded outbreak as:

$$onset_{1t} = \begin{cases} 0, & \text{if conflict}_t = 0 \\ 1, & \text{if conflict}_t = 1\& \text{ conflict}_{t-1} = 0 \\ 1, & \text{if conflict}_t = 1\& \text{ conflict}_{t-1} = 1 \end{cases} \tag{1.18}$$

then conflict outbreak and conflict incidence would exactly equal one another; conflict incidence can therefore be thought of as a measure of civil conflict onset which treats every conflict period as independent. Results from the previous section have already presented evidence of statistically significant level and uncertainty effects for conflict incidence.

On the other hand, conflict outbreak might also be measured as:

$$onset_{2t} = \begin{cases} 0, & \text{if conflict}_t = 0 \\ 1, & \text{if conflict}_t = 1\& \text{ conflict}_{t-1} = 0 \cdot \\ 0, & \text{if conflict}_t = 1\& \text{ conflict}_{t-1} = 1 \end{cases} \tag{1.19}$$

This measure treats all periods of conflict incidence immediately following another period of conflict incidence as a continuation of the former, strictly capturing transitions from states of non-conflict to conflict. Table 1.11 replicates the estimation of the models in Table 1.10, showing the results are robust to incorporating this modified measure of conflict outbreak as the dependent variable. Across all columns, an increase in the change in the net export commodity price index leads to a lower likelihood of transition from peace to conflict. This first-moment effect is significant at $\alpha = 0.01$. An increase in the change in uncertainty of net export commodity prices results in a higher likelihood of transition from non-conflict to conflict. This second-moment effect is significant at typical statistical significance levels.

	IV-LPM-FE	LOGIT-FE	PROBIT-RE	IV-LPM-FE	LOGIT-FE	PROBIT-RE
ΔPI_t	-0.022**	-2.906**	-0.954**	-0.021**	-2.922**	-0.924**
	(0.008)	(0.963)	(0.297)	(0.008)	(0.969)	(0.302)
ΔVarPI_t	0.042+	5.889*	2.099+			
	(0.023)	(2.746)	(1.076)			
$\Delta \text{Vol}_1\text{PI}_t$				0.036+	4.823**	1.726*
				(0.020)	(1.557)	(0.824)
Δincome_t	-0.865	-5.548+	-2.884*	-0.860	-5.519+	-2.848*
	(0.991)	(2.863)	(1.363)	(0.995)	(2.823)	(1.353)
Constant			-2.539***			-2.536***
			(0.044)			(0.044)
$\ln \sigma_u^2$			-13.825***			-13.858**
			(3.816)			(4.226)
N	9108	6148	10385	9108	6148	10385
Clusters	73	44	75	73	44	75
Overall p	0.034	0.002	0.001	0.037	0.001	0.001
Over-i.d. test p	0.505			0.475		
Endog. test p	0.406			0.416		

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.11: Effect of price uncertainty (variance and s.d.) on modified onset measure

Different Index Weights

In the results above, price and volatility indices were constructed using time-invariant weights calculated as net export shares of individual commodities from the year 2006. The use of time-invariant weights is partly motivated by data limitations which preclude the use of time-varying weights throughout the entire sample period. However, under the assumption that the agricultural commodities which a developing country produces in a given year maintain a relatively stable role in the country's net export basket vis-à-vis the other agricultural commodities, the findings will remain valid. For example, this entails that a country that is primarily a net exporter of wheat in one year does not switch its primary net export commodity to bananas in another year of the sample period. Nevertheless, to check whether the effect of commodity price volatility on the incidence and outbreak of conflict is sensitive to the choice of reference year used for the weights, I recalculate price and volatility indices using net export shares from the year 1996 to weigh the individual commodities.

Tables 1.12 and 1.13 illustrate that despite the drop in the number of countries included in the estimation in comparison to the baseline specification, there is evidence

for price level and uncertainty effects on both the incidence and outbreak of conflict when using export weights from an earlier year. Table 1.12 shows a sample of results for the effect of price uncertainty and various measures of income on conflict incidence. Across all specifications, increases in the change in the net export commodity price index are associated with decreased likelihood of conflict incidence, while a greater increase in price uncertainty over the previous quarter is associated with an increased likelihood of conflict incidence, as predicted by the model. The price level effect is only significant when controlling for international reserves, but the uncertainty effect is significant in all columns of Table 1.12. Results are qualitatively similar when examining conflict outbreak, as seen in Table 1.13. The level and volatility effects are significant and of the correct sign in all columns. The instrument sets for the IV specifications in Tables 1.12 and 1.13 are generally the same as for the corresponding specifications which use year 2006 shares as weights, with the exception of the first column of Table 1.13, which uses the fourth through the eighth lags of changes in rainfall to instrument for changes in GDP per capita.

	LOGIT-RE	LOGIT-FE	IV-LPM-FE	LOGIT-RE	IV-LPM-FE	LOGIT-RE
ΔPI_t	-0.458	-0.459	-0.058*	-0.605*	-0.038	-0.401
	(0.299)	(0.298)	(0.029)	(0.271)	(0.026)	(0.295)
ΔVarPI_t	1.294**	1.295**	0.109*	1.218**	0.087+	1.254*
	(0.423)	(0.421)	(0.051)	(0.438)	(0.049)	(0.505)
Δrain_t	0.036	0.036				
	(0.022)	(0.022)				
$\Delta \text{reserves}_t$			-0.033	-0.074		
			(0.032)	(0.167)		
Δincome_t					-0.755	-2.942
					(3.322)	(2.910)
Constant	-2.959***			-3.521***		-2.875***
	(0.521)			(0.677)		(0.764)
$\ln \sigma_u^2$	2.369***			2.747***		2.946***
	(0.275)			(0.341)		(0.442)
N	8672	6594	6677	7567	6073	7017
Clusters	52	38	48	52	49	51
Overall p	0.026	0.028	0.130	0.024	0.255	0.067
Over-i.d. test p			0.463		0.174	
Endog. test p			0.346		0.530	

$+ \ p < 0.10$, $^* \ p < 0.05$, $^{**} \ p < 0.01$, $^{***} \ p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.12: Price volatility and conflict incidence, using year 1996 net export shares as weights

	IV-LPM-FE	LPM-RE	LPM-FE	LOGIT-RE	LOGIT-FE	PROBIT-RE
ΔPI_t	-0.035*	-0.030*	-0.031*	-2.468**	-3.356**	-1.163**
	(0.016)	(0.013)	(0.013)	(0.756)	(0.945)	(0.380)
$\Delta \text{Vol}_1 \text{PI}_t$	0.066+	0.060*	0.062*	3.598*	6.324*	1.934*
	(0.038)	(0.030)	(0.031)	(1.677)	(2.665)	(0.848)
Δincome_t	-1.535	-0.133	-0.124	-13.629+	-11.283	-4.609
	(7.006)	(0.096)	(0.100)	(7.756)	(8.987)	(3.013)
Constant		0.008***	0.008***	-4.921***		-2.444***
		(0.001)	(0.000)	(0.172)		(0.060)
$\ln \sigma_u^2$				-12.480***		-13.837***
				(3.193)		(1.097)
N	4512	5338	5338	5338	3170	5338
Clusters	44	47	47	47	31	47
Overall p	0.218	0.028	0.068	0.004	0.003	0.010
Over-i.d. test p	0.168	0.290				
Endog. test p	0.748					

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 1.13: Price volatility and conflict onset, using year 1996 net export shares as weights

1.5 Conclusion

The contribution of this paper is to analyze the effects of commodity price fluctuations in generating political conflict in developing countries. I find that in addition to price level shocks, increased uncertainty about future net export commodity prices is a significant predictor of conflict incidence in developing countries. I also present evidence that net export commodity price uncertainty affects the outbreak of conflict as well as its incidence. These findings have several implications for development policy. The first is that restrictions on export quantities, export licenses, subsidies or other state intervention in the market that reduce the volatility of relative prices faced by citizens in a developing country may generate a more conducive environment for a new or transitioning government to consolidate its role. Only a model which fully specifies the extent of welfare loss from both political instability and the economic inefficiency entailed by these price controls can reveal how a policymaker should optimally react to these distortions. It may be that for a newly-formed government, an active state role in the economy is desirable, while economic liberalization remains important for long-run growth.

A second implication is that financial market regulation in the developed world can contribute to political stability, or the lack thereof, in less developed countries. In the

past decade, commodity markets have become increasingly dominated by financial speculators, and the introduction of tools such as commodity index funds, high frequency and algorithmic trading, and deregulated over-the-counter trading have substantially increased the volatility of prices faced by food producers.[7] In 2012, U.N. special rapporteur on the right to food Olivier De Schutter remarked, "What we are seeing now is that these financial markets have developed massively with the arrival of these new financial investors, who are purely interested in the short-term monetary gain and are not really interested in the physical thing – they never actually buy the ton of wheat or maize; they only buy a promise to buy or to sell. The result of this financialisation of the commodities market is that the prices of the products respond increasingly to a purely speculative logic. This explains why in very short periods of time we see prices spiking or bubbles exploding, because prices are less and less determined by the real match between supply and demand." The years after the 2008 global financial crisis have seen an increased effort to expand the government's role in commodity markets in order to curtail such activities, such as the "Over-the-Counter Derivatives Markets Act of 2009" drafted by the U.S. Treasury Department, H.R. 4173 (the "Wall Street Reform and Consumer Protection Act of 2009"), and provisions in the 2010 Dodd-Frank financial reform law. The discussion in this paper suggests that the enlargement of government intervention in commodity markets, in addition to affecting the incentives and behavior of financial market participants, may have an impact on the political process of developing countries.

[7]Descriptive statistics of these recent changes in commodity markets and their impact on food prices is presented in the World Development Movement report *Broken Markets*, available at http://www.wdm.org.uk/sites/default/files/Broken-markets.pdf.

Appendix A

Data and Sources

Data	Source
Annual per-capita GDP series	World Bank, World Development Indicators: *http://data.worldbank .org/indicator/NY.GDP.PCAP.CD*
Annual commodity import and export values	United Nations Commodity Trade Statistics Database: *http://data.un. org/Explorer.aspx?d=ComTrade*
Military expenditure and personnel	Correlates of War Project: National Material Capabilities v4.0 *http://correlatesofwar.org/*
Monthly global commodity prices Year 2000 constant price	World Bank GEM Commodities *http://data.worldbank.org/data-catalog/commodity-price-data*
Monthly global rainfall data (unadjusted series), station avg.	Global Historical Climatology Network v2 *http://www.ncdc.noaa.gov/ghcnm/v2.php*
Nonviolent and Violent Conflict Outcomes (NAVCO) Dataset	*http://echenoweth.faculty.wesleyan.edu/ research-and-data/*
UCDP/PRIO Armed Conflict Dataset v.4-2010, 1946-2009	*http://www.pcr.uu.se/research/ucdp/ datasets/ucdp_prio_armed_conflict_dataset/*
Rating of political rights and civil liberties: Combined Average Ratings - Independent Countries	Freedom House: Freedom in the World Index 2011: *http://www.freedomhouse.org/ template.cfm?page=25& year=2011*
International Reserves (measured in billions of SDR)	IMF *International Financial Statistics* Quarterly
Household consumption expenditure	UNdata: *http://data.un.org.*

Data	Source
(including non-profit institutions serving households) in constant 2005 U.S. dollars	
Population	UNdata: *http://data.un.org.*
GDP per capita: OECD total	OECD Statistics

Appendix B

Relative import prices

Although the focus of this paper has been to measure the effect of fluctuations in the level and uncertainty of relative export prices on political conflict, the model in Section 1.3 can be easily recast to yield similar conclusions regarding relative import prices. By rewriting the citizens' budget constraint as: $\tilde{P}_t C_t = (1-\tau)m$, where $\tilde{P}_t$ is the exogenous price of imports in units of the exported commodity endowment, and assuming that $\tilde{P}_t$ follows a log-normal distribution, the model can be solved in exactly the same way as the model in the paper. Citizens' intertemporal utility at the end of period 1 is:

$$\begin{aligned} U_1 &= \log \tfrac{(1-\tau)m}{\tilde{P}_1} + \beta E_1 \left[\log \tfrac{(1-\tau)m}{\tilde{P}_2}\right] \\ &= \log \tfrac{(1-\tau)m}{\tilde{P}_1} + \beta \left[\log (1-\tau)m - \mu\right] \\ &= \log \tfrac{(1-\tau)m}{\tilde{P}_1} + \beta \left[\log (1-\tau)m + \log E_1[\tfrac{1}{\tilde{P}_2}] - \tfrac{\sigma^2}{2}\right] \end{aligned} \tag{B.1}$$

where the last two lines in (B.1) follow from the properties of the log-normal distribution. Suppose there is a permanent, positive shock to the mean of the log price distribution at the end of period 1, which could happen in the realistic case that the log of prices is non-stationary. By the second line of (B.1), the citizens' end-of-period expected utility is lowered, making it more likely the citizens will not achieve their reservation utility level and, consequently, overthrow the dictator. Suppose in period 1, there is a positive, mean-preserving, permanent shock to the variance of the log price distribution (i.e. $\tilde{P}$ is held constant). By the third line of (B.1), this shock to uncertainty also makes it more likely that revolution will occur. This analysis suggests that both first-moment and second-moment relative import price shocks should play a role in determining conflict

in developing countries. The only difference relative to the model in the paper is that when looking at relative import prices, positive rather than negative shocks to price level lower utility and make conflict more likely to occur. The political effect of uncertainty regarding the relative price distribution remains the same in the model in Section 1.3.

The empirical evidence for global commodity import prices affecting conflict is far weaker than for the export side. I generate net import-share weighted commodity price level (PI_{imp}) and uncertainty (VarPI_{imp}) indices following the same method described in Section 1.4. Summary statistics given in Table B.1 indicate no obvious difference in variation between the net import and net export weighted indices that would yield a drastically different magnitude of precision in the estimation of the models. However, the results shown in Table B.2 for the estimated effect of a change in the net import share-weighted commodity price level and uncertainty index on the incidence of conflict give a different story. Although positive changes in commodity net import price levels and uncertainty are associated with higher risk of civil conflict as predicted by the model, the estimated coefficients are not significant. There is a consistent lack of significance for the net import share-weighted indices for all the analogous specifications of the models estimated in the Results section of this paper. The failure to find any statistically significant relationship between net import-weighted global commodity price fluctuations and civil conflict may be due to the existence of price controls for imported consumption goods in poorer countries. Domestic price controls can make global prices less relevant for welfare (and, consequently, political stability) in a country. Governments of lower-income, small open economies may find it easier to impose price controls on consumption goods than to dictate the prices for which they can sell domestically produced commodities abroad, explaining the asymmetric results between export and import prices on conflict. I leave a detailed investigation of this asymmetric effect as an issue for further research and argue that the results presented in the body of the paper still indicate that fluctuations in global commodity price levels and uncertainty can impact political stability in developing countries.

Variable	Obs	Mean	Std. Dev.	Min	Max
PI_{imp}	16544	0.836	0.280	0.184	3.556
$VarPI_{imp}$	16544	0.057	0.067	0.000	1.189

Table B.1: Summary statistics for net import share weighted price and uncertainty series, quarterly frequency

	IV-LPM-FE	LPM-FE	LOGIT-RE	LOGIT-FE	PROBIT-RE
$\Delta PI_{imp,t}$	0.004	0.010	0.116	0.116	0.076
	(0.030)	(0.027)	(0.317)	(0.315)	(0.171)
$\Delta VarPI_{imp,t}$	0.023	0.014	0.193	0.188	0.095
	(0.061)	(0.054)	(0.632)	(0.627)	(0.344)
$\Delta income_t$	-0.697	-0.299	-2.506	-2.455	-1.330
	(2.322)	(0.192)	(1.814)	(1.781)	(0.910)
Constant		0.224***	-3.028***		-1.660***
		(0.000)	(0.532)		(0.304)
$\ln \sigma_u^2$			2.717***		1.359***
			(0.311)		(0.327)
N	10226	11829	11829	7512	11829
Clusters	84	87	87	56	87
Overall p	0.667	0.231	0.297	0.304	0.277
Over-i.d. test p	0.209				
Endog. test p	0.969				

$+ p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table B.2: Effect of net import-weighted price level, volatility, and income on conflict incidence

Appendix C

Countries included in the dataset

The countries included in the dataset are:

Afghanistan, Angola, Azerbaijan, Bangladesh, Belize, Benin, Bhutan, Bolivia, Bosnia and Herzegovina, Botswana, Bulgaria, Burkina Faso, Burundi, Cambodia, Cameroon, Chad, Chile, China, Colombia, Congo, Dem. Rep., Congo, Rep., Costa Rica, Cote D'Ivoire, Croatia, Cuba, Curacao, Czech Republic, Djibouti, Dominican Republic, Ecuador, Egypt, El Salvador, Equatorial Guinea, Eritrea, Ethiopia, Fiji, The Gambia, Ghana, Grenada, Guatemala, Guinea, Guinea-Bissau, Guyana, Haiti, Honduras, India, Indonesia, Jamaica, Kazakhstan, Kenya, Kiribati, Korea, Dem. Rep., Kosovo, Lao PDR, Lebanon, Lesotho, Liberia, Macedonia, Madagascar, Malawi, Malaysia, Mali, Marshall Islands, Mauritius, Mexico, Micronesia, Fed. Sts., Moldova, Mongolia, Mozambique, Myanmar, Namibia, Nepal, Nicaragua, Niger, Nigeria, Northern Mariana Islands, Pakistan, Papua New Guinea, Paraguay, Peru, Philippines, Puerto Rico, Rwanda, Samoa, Sao Tome and Principe, Senegal, Serbia, Sierra Leone, Sint Maarten (Dutch part), Slovak Republic, Solomon Islands, Somalia, South Africa, Sri Lanka, St. Lucia, St. Martin (French part), St. Vincent and the Grenadines, Swaziland, Syria, Tajikistan, Tanzania, Thailand, Timor-Leste, Togo, Tunisia, Turkey, Turkmenistan, Uganda, Ukraine, Uzbekistan, Vanuatu, Vietnam, West Bank and Gaza, Yemen, Rep., Zambia, Zimbabwe

Depending on data availability and estimation method, only a subset of the full list may be incorporated into any particular regression presented in the paper. As discussed in Section 1.4, countries listed above are included on the basis of their World

Bank classification as low- and lower-middle income countries in 1987. This definition encompasses countries which may not be appropriately considered small open economies today, particularly China and India. However, it can be argued due to political factors, these countries resembled closed economies for a large part of the sample period and their inclusion may actually bias the results against finding a relationship between global price fluctuations and internal political conditions. For example, in India:

> Before the process of reform began in 1991, the government attempted to close the Indian economy to the outside world. The Indian currency, the rupee, was inconvertible and high tariffs and import licensing prevented foreign goods reaching the market. India also operated a system of central planning for the economy, in which firms required licenses to invest and develop....
>
> The central pillar of the policy was import substitution, the belief that India needed to rely on internal markets for development, not international trade - a belief generated by a mixture of socialism and the experience of colonial exploitation. Planning and the state, rather than markets, would determine how much investment was needed in which sectors. ("India: the economy," BBC News, 3 December 1998)

The econometric results presented in the paper above are robust to the inclusion of China and India in the sample. Table C.1 below shows a sampling of the results after dropping China and India from the dataset.

	LOGIT-RE	LOGIT-FE	IV-LPM-FE	PROBIT-RE	IV-LPM-FE	PROBIT-RE
	(incidence)	(incidence)	(incidence)	(incidence)	(onset)	(onset)
ΔPI_t	-0.367+	-0.368+	-0.050*	-0.321**	-0.030**	-0.971**
	(0.207)	(0.207)	(0.021)	(0.118)	(0.011)	(0.299)
ΔVarPI_t	0.931**	0.931**	0.082*	0.500*	0.055+	2.223+
	(0.334)	(0.333)	(0.040)	(0.236)	(0.029)	(1.135)
Δrain_t	0.018	0.018				
	(0.015)	(0.015)				
$\Delta\text{reserves}_t$			-0.119	0.076		
			(0.100)	(0.068)		
Δincome_t					-0.845	-3.708+
					(1.282)	(1.940)
Constant	-2.968***			-1.916***		-2.472***
	(0.405)			(0.271)		(0.057)
$\ln \sigma_u^2$	2.144***			1.091***		-3.981
	(0.241)			(0.305)		(9.280)
N	12249.000	9208.000	9099.000	10128.000	6865.000	7941.000
Clusters	72.000	53.000	67.000	71.000	66.000	70.000
Overall p	0.012	0.014	0.034	0.004	0.072	0.001
Over-i.d. test p			0.741		0.493	
Endog. test p			0.067		0.606	

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table C.1: Price volatility and conflict likelihood, omitting China and India

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Chapter 2

New perspectives on corruption contagion: evidence from Asia, Latin America and the Middle East

2.1 Introduction

Is corruption within one country affected by corruption within another? Understanding the interactions between political-economic culture across countries can allow us to better grasp the implications of greater global and regional integration in recent history. Until now, few studies have examined this question in detail due to the difficulty of measuring corruption and paucity of consistent data over an adequate time span. In this paper, I use a panel dataset of countries in Asia, Latin America, and the Middle East over a span of fifteen years to examine how domestic corruption reacts to the culture of corruption in the region in which the country is located. Contrary to the results of past literature, I find evidence that a reduction in regional corruption can actually lead to a worsening of corruption within a country, and vice versa. If, in an open economy, regional graft lowers the level of income that a rent-seeking government can tax, a reduction in regional corruption can increase the marginal benefit (from

the perspective of the government) of imposing a more extractive domestic policy by increasing the pool of exploitable funds. Though economists and political scientists have emphasized the growth in trade, communications, and financial transactions across borders as powerful agents of cultural change, my results offer an economic reason for why national corruption will be an enduring institution in a more interconnected world.

There are several reasons that corruption could be linked across countries. Previous studies have established a relationship between corruption and levels of foreign direct investment in developing countries, and found that FDI flows to one country can be affected by flows of FDI to other countries. If countries compete for FDI, strategic pressures could motivate governments to reduce levels of graft in their country to become relatively more appealing than their peers as investment destinations. Alternatively, a country's corruption could worsen even as corruption amongst its regional peers declined if reduced regional corruption increases the incentives for a government to engage in more graft domestically. A rent-seeking government may find that lower regional corruption has positive externalities for domestic income: foreign investors with an imperfect capability to monitor idiosyncratic corruption may choose to divert more capital to countries located in a region with lower overall corruption, or citizens who consume traded goods may have more incentive to substitute from leisure into consumption and labor as the corruption tax on traded goods declines. This could increase the pool of funds that are exploitable by a rent-seeking government and provide incentive to engage in more corrupt behavior even while other countries in the region engage in less. I build a model where external corruption affects agents' choice of labor supply and consequently, the domestic government's optimal rent extraction rate. The model yields the result that corruption within countries moves in the opposite direction as external corruption.

Anecdotal evidence of countries located in the same geographic region can be found to support this hypothesis. Global Integrity, an international nonprofit organization dedicated to documenting trends in corruption and governance, has released annual studies since 2006 of the effectiveness of anti-corruption mechanisms in various countries. Key findings in its 2009 Global Integrity Report include Mongolia and Vietnam being

placed on a watchlist of countries highly susceptible to corruption even as China, with whom both countries share a border, was removed from that same list. In Eastern Europe, Ukraine was added to the watchlist while Georgia was removed. In Sub-Saharan Africa in 2008, Nigeria was found to have improved its anti-corruption framework from the previous year while the efficacy of anti-corruption laws in Ethiopia, Ghana and Kenya significantly declined.[1] Although these individual case studies may suggest a more pervasive trend, it is unclear whether they are representative of countries' political experiences overall. In order to see how corrupt behavior is correlated internationally, it is necessary to harmonize economic theory with observed patterns in the data. In this paper I will present statistical evidence for negative cross-country corruption linkage.

The structure of this paper is as follows: In Section 2.2, I discuss previous work related to the economics of corruption; Section 2.3 presents a simple model of cross-country corruption interaction; Section 2.4 describes the data and offers econometric evidence supporting the key results of the model, and Section 2.5 concludes.

2.2 Review of Literature

Economists have long sought to understand why corruption arises and how incentives might be created to attenuate the problem. In an early contribution, Rose-Ackerman considers how market structure influences private incentives to pay bribes to procure government contracts, finding greater scope for bribery when government is the sole purchaser of a good in comparison to a situation where a good is also available on the private market. Rose-Ackerman also analyzes the effectiveness of penalties on government officials and firms that engage in corrupt behavior, finding that under some conditions, they may be ineffective in deterring bribery (Rose-Ackerman 1975). Shleifer and Vishny demonstrate that corruption can be exacerbated if government agencies act independently in extorting bribes when providing a government good. If entry into bribe-collection is free, the total bribe can rise to infinity and no governemnt good will be provided (Shleifer and Vishny 1993). Marcouiller and Young show that a predatory

[1] Annual country reports and summaries of key findings from the Global Integrity Report are available at `http://www.globalintegrity.org/report`.

government that increases taxation and graft at the expense of a shrinking formal economy may be acting rationally, and that the optimal level of government predation is determined by the elasticity of substitution between the goods produced in the formal and informal sectors of the economy (Marcouiller and Young 1995). Bliss and Di Tella argue that the number of firms in an economy cannot be used as an exogenous measure of competition when considering the relationship between competition and corruption, as corruption may affect the extent of competition. When considering deeper measures of competition, such as similarity of production costs and a tendency of lower overhead costs relative to profits, the authors find that increases in competition may not necessarily lower corruption (Bliss and Di Tella 1997). My analysis will differ from these papers in that I consider the spillover effects of international corruption on domestic corruption; additionally, I will consider a broader definition of corruption that includes transactions in which one party may not be a willing participant.

My paper is more closely related to empirical work quantifying the impact of corruption. Mauro shows evidence from a cross-section of countries that higher corruption is linked to lower investment, thereby lowering growth (Mauro 1995). Wei analyzes flows of foreign direct investment from fourteen source countries to 45 host countries, finding that increases in the corruption level in host governments reduce inward FDI. He also finds that investors in the United States are not more averse to corruption than investors in other countries despite the Foreign Corrupt Practices Act of 1977 (Wei 2000). In a related paper, he demonstrates that domestic crony capitalism, defined as an economic environment where commercial disputes are resolved and resources allocated in favor of friends and relatives of those in power, is associated with a higher external loan-to-FDI ratio, increasing the likelihood of currency crises. Because direct investors are more likely to be exposed to local officials and bureaucracy, higher corruption increases the cost of equity participation and causes investors to favor debt contracts, including bank loans, as a less risky mode of investment (Wei 2001). Javorcik and Wei use firm-level data on FDI in Eastern Europe and the former Soviet Union in the 1990s to analyze the impact of corruption on the choice of entry mode, as well as the volume of FDI. They find that corruption reduces inward FDI, corroborating previous studies, and that higher

corruption is associated with more joint ventures as local partners who are familiar with bureaucratic procedures become more valuable. However, with sufficient technological sophistication, foreign investors prefer wholly-owned projects to joint venture firms; the authors find that American investors were more reluctant to form joint ventures in more corrupt countries (Javorcik and Wei 2009).

In contrast to the previous papers which consider corruption as an independent variable and measure its effect on other outcomes, I will consider corruption itself as the variable of interest. Ades and Di Tella present the first empirical study of the causes of corruption, using cross-sectional data from the 1980s and 1990s to show that countries where firms enjoy higher rents (implying less competition between firms or ineffective antitrust measures) tend to have higher corruption levels (Ades and Di Tella 1999). In a similar analysis, Treisman finds that countries with Protestant traditions, histories of British rule, more developed economies, long exposure to democracy and, less robustly, higher imports had lower levels of corruption while countries with federal governments tended to be more corrupt (Treisman 2000). Emerson shows a negative correlation between corruption and the level of competition in a country, and finds evidence from international cross-sectional data that higher levels of education and more political rights and participation decrease corruption (Emerson 2006). These authors do not consider, however, the effect of corruption in neighboring countries on domestic corruption.

This paper is most closely related to work that seeks to measure the contagion effects of corruption. Sandholtz and Gray argue that higher international integration leads to lower corruption and present evidence that greater participation in international organizations, higher international economic integration, economic development, and democracy reduce domestic corruption perception levels (Sandholtz and Gray 2003). To assess the possibility that cultural similarities provide a channel for transmission of norms related to corruption, Sandholtz and Gray also include a measure of average perceived corruption levels in bordering countries, finding a positive correlation with domestic perceived corruption when OECD countries are included in the sample. Goel and Nelson use state-level U.S. data averaged over the years 1995-2004 to show that a 10% increase in corruption in neighboring states increases the number of convictions

due to corruption within a state by 4-11%. The authors also find some evidence tying corruption to economic performance, noticing that federal public corruption convictions increase with the unemployment rate and decrease with gross state product, as well as evidence of a contagion effect for murder rates (Goel and Nelson 2007). Becker et al. use a cross-section of 123 economies to show that a higher level of perceived corruption in adjacent economies leads to domestic increases in perceived corruption (Becker, Egger and Seidel 2009). Corruption spillover effects are also present when geographical distance (rather than common borders) and common official language use are utilized to define a particular country's neighbors, but not when trade weights are used. In contrast to Becker et al., I will show that considering trade-weighted measures of corruption in a panel context lead to a reverse contagion effect.

In the next section, I present a simple model that describes a spillover relationship of international corruption that is very different from the contagion effect which has been depicted in the literature. The conclusions of this model will rationalize the econometric exercise in Section 2.4.

2.3 A Simple Model

2.3.1 Model Setup

In this section I construct a simple model of international corruption interaction from a partial equilibrium, small open economy perspective. The advantages of this model are that it highlights an interrelationship between country-level and region-level corruption that differs from the relationship found in past literature, and it can serve as a useful starting point to motivate the empirical analysis in Section 2.4.

Households

Consider a small open economy ruled by a corrupt, rent-seeking government and populated by a large number of identical households with preferences described by the following utility function:

$$U(c, L) = c + \ln(1 - L) \tag{2.1}$$

where c denotes consumption and L denotes labor effort. Households gain linear utility from consumption and log utility from leisure, where households' time endowment is normalized to unity.

Households engage in production of a good that is not domestically consumed, but which is valued by consumers in the rest of the region. The production technology is $F(L) = WL$, where W is the marginal product of labor which I henceforth refer to as the wage. Labor income (production) is subject to graft imposed by corrupt government officials which operates as a tax. Income net of the corruption tax τ is used to import the consumption good from abroad at price P^*. Exporting countries in the rest of the region also suffer from corruption which affects domestic consumers in the form of an ad valorem corruption tariff. The budget constraint that faces households can thus be written as:

$$(1-\tau)WL \geq (1+\tau^*)P^*c \tag{2.2}$$

Forming the Lagrangian,

$$\mathcal{L}(c, L, \lambda) = c + \ln(1-L) + \lambda\left((1-\tau)WL - (1+\tau^*)P^*c\right) \tag{2.3}$$

the first-order conditions are:

$$1 - \lambda(1+\tau^*)P^* = 0 \tag{2.4}$$

$$-\frac{1}{1-L} + \lambda(1-\tau)W = 0 \tag{2.5}$$

$$(1-\tau)WL = (1+\tau^*)P^*c \tag{2.6}$$

Equations (2.4) and (2.5) define the marginal rate of substitution between consumption and labor as:

$$\frac{\frac{\partial U}{\partial c}}{\frac{\partial U}{\partial L}} = -\frac{(1+\tau^*)P^*}{(1-\tau)W} \tag{2.7}$$

For a given level of wages and prices, as corruption in the rest of the region τ^* increases, fewer hours must be worked (i.e. more leisure must be consumed) in order to compensate for a one-unit marginal reduction in consumption. Higher regional corruption can

increase the effective price of imported goods as consumers are forced to pay a corruption premium. The resulting increase in disutility then causes the citizen to substitute from consumption into leisure and earn lower income. The labor-leisure tradeoff channel generated by movements in foreign corruption can affect the decisions of domestic officials who extract corruption rents from labor income. Equations (2.4)–(2.6) can be combined to form the household's labor supply equation:

$$L = 1 - \frac{(1+\tau^*)P^*}{(1-\tau)W} \tag{2.8}$$

which also shows that increases in foreign corruption reduce domestic labor supply.

Government

The domestic government chooses a rate of graft $\tau \in (0,1)$ to maximize its rents:

$$\max_{\tau} \tau W L \tag{2.9}$$

Substituting for L with the household labor supply function, this can be rewritten as:

$$\max_{\tau} \tau W \left(1 - \frac{(1+\tau^*)P^*}{(1-\tau)W}\right) \tag{2.10}$$

As the second-order condition for this optimization problem is globally satisfied,[2], taking the derivative of (2.10) with respect to τ and setting equal to zero yields two possible solutions for the extraction rate τ that maximizes government rents:

$$\tau = \frac{W + [P^*W(1+\tau^*)]^{\frac{1}{2}}}{W} \quad \text{or} \quad \tau = \frac{W - [P^*W(1+\tau^*)]^{\frac{1}{2}}}{W} \tag{2.11}$$

The first root is outside the range for feasible values of τ and is ignored. The domestic government thus sets its rate of corruption as:

$$\tilde{\tau} = \max\left\{0, \frac{W - [P^*W(1+\tau^*)]^{\frac{1}{2}}}{W}\right\} \tag{2.12}$$

[2]The second derivative of the government objective function is $\frac{2P^*\tau(1+\tau^*)}{(\tau-1)^3} - \frac{2P^*(1+\tau^*)}{(\tau-1)^2} < 0$.

Equation (2.12) shows that domestic corruption is linked to consumer prices, wages (or income), as well as corruption in the rest of the region.

The key result of this paper lies in the manner in which domestic corruption responds to changes in corruption in the rest of the region. This can be seen by taking the derivative of (2.12) with respect to τ^*:

$$\frac{\partial \tilde{\tau}}{\partial \tau^*} = \begin{cases} 0 & \text{if } \tilde{\tau} = 0 \\ \dfrac{-P^*}{2\left[P^*W(1+\tau^*)\right]^{\frac{1}{2}}} < 0 & \text{if } \tilde{\tau} = \dfrac{W - \left[P^*W(1+\tau^*)\right]^{\frac{1}{2}}}{W} \end{cases} \tag{2.13}$$

For a small change in corruption in the rest of the region, either there is no contagion or an increase (decrease) in regional corruption leads to a decrease (increase) in corruption within country i. In the latter case, the reduction of corruption in the neighboring region effectively increases the income of a citizen within a given country, enlarging the pool of exploitable funds and giving incentive for a rent-seeking government to engage in more extractive policies. This result contrasts with the intuition in other papers that corruption is contagious, or that corruption in one country tends to move in the same direction as corruption endemic to neighboring countries in the region. This finding is also not specific to this modeling approach; Persson and Tabellini build a two-country general equilibrium model of tax competition that features international capital mobility and governments which maximize consumer utility. They find that the policy response of a country to the capital taxation rate of another country is ambiguous and dependent on the concavity of the utility function (Persson and Tabellini 1995). In their paper they assume that the utility function is not very concave, allowing them to draw upward sloping best-response curves, but this need not be true. If corruption can be thought of as a type of tax rate on consumers, the same model can be used to justify the corruption substitutability result discussed here. In Appendix D, I describe a framework for a general equilibrium model of corruption that also exhibits this strategic substitutability.

2.4 Empirical Analysis

In this section, I evaluate the primary hypothesis from the model described above, namely that domestic corruption is negatively related to corruption in the rest of the region in which a country is located.

2.4.1 Data and Methodology

Data on corruption are taken from the Heritage Foundation's 'Freedom from corruption' index, a component of the Index of Economic Freedom released annually since 1995. 'Freedom from corruption' is measured using a scale from 0 to 100, where 100 represents the maximum possible level of freedom from corruption. The index draws upon polls of experts as well as individual surveys of members of the business community and the local population to measure perceived corruption within a country.[3] Although perceived corruption indices are subjective and an imperfect substitute for the elusive ideal of an objective, direct measure of corruption, Treisman notes several reasons that make the usage of such indices informative: high correlation over time, which suggests that there is relatively small risk of outlying inputs skewing the ratings in a particular year; high correlation with indices of corruption compiled by other organizations using different methodologies; and the prevalence of studies showing that perceived corruption ratings can predict various measures of economic performance, suggesting that perceived corruption may be as important as corruption itself (Treisman 2000). Additionally, attempts to objectively measure absolute levels of corruption through alternative metrics such as the size of bribe payments, number of mentions of corruption in the national media, or number of corruption-related court cases brought against public officials may in fact measure other sociopolitical characteristics of the country such as extent of freedom of the press or efficacy of legal prosecutors.

A possible concern that may emerge from the use of indices based on perception is whether they are comparable over time. If perceptions of corruption are based on

[3]Further information on the methodology used in the calculation of the Heritage Foundation's freedom from corruption index can be found at `http://thf_media.s3.amazonaws.com/index/pdf/2011/Index2011_Methodology.pdf`. Methodology and sources for Transparency International's Corruption Perceptions Index, which forms the basis for the Freedom from corruption index in later years, can be found at `http://cpi.transparency.org/cpi2011/in_detail/`.

a relativistic evaluation (i.e. how corrupt a country is relative to the environment of corruption in other countries), a numerical score in one year may have different implications than the exact same score in another year. This might occur, for instance, if a region or the entire world experiences a simultaneous shift in perceived corruption. However, to the extent that all countries within a region experience a simultaneous change (either an increase or a decrease) in perceived corruption within a certain period, this will be captured by the inclusion of time fixed effects. In practice, high correlation of country-level corruption indices over time make such situations unlikely (see Table 2.1). Furthermore, in the econometric exercise in the next section I demonstrate that the presence of negative corruption spillovers is robust to several specifications of the time fixed effects. Summary statistics for countries' freedom from corruption indices, partitioned by region, are presented in Table 2.1.

Region	Obs	Mean	Std. Dev.	Min	Max	Corr. with first lag
Asia	323	35.78328	23.57641	4	94	0.9702
Latin America	319	34.55486	14.78513	10	79	0.9174
Middle East	220	45.28182	24.10693	10	90	0.9613

Table 2.1: Summary statistics for freedom from corruption index by region

For each country, I generate a measure of corruption in the rest of the region in which the country is located. As the model suggests that economic openness (via trade) is the primary channel through which domestic income is related to regional corruption, I construct an openness-weighted aggregate of the 'Freedom from corruption' indices of each country's regional neighbors. The weight that a country's regional peer receives in this index is calculated as:

$$\omega_{ijt} = \frac{openness_{jt}}{\sum_{k\setminus\{i\}} openness_{kt}} \tag{2.14}$$

ω_{ijt} is the weight of country j in country i's regional 'Freedom from corruption' index at time t. In the denominator are all countries k in the region in which country i is located, not including country i itself; as country i does not appear in its own 'rest of region' index, the weights associated with the other countries sum up to unity. The resultant index is specific to each country and can be used to measure the effect of

aggregate perceived corruption in a country's regional neighbors on domestic levels of perceived corruption. For robustness, I consider several measures of openness: trade (defined as the sum of exports and imports of goods and services) as a share of GDP, exports as a share of GDP, and imports as a share of GDP. Data on these variables are from the World Bank's *World Development Indicators* and are available until 2010. Countries included in the sample and the region in which they are located are described in Table 2.2. In Figures 2.1 to 2.3, I plot the country average freedom from corruption index versus the regional average index over the sample period. The figures suggest a negative cross-sectional correlation between a country's perceived corruption and the trade-weighted average of that of its regional peers.

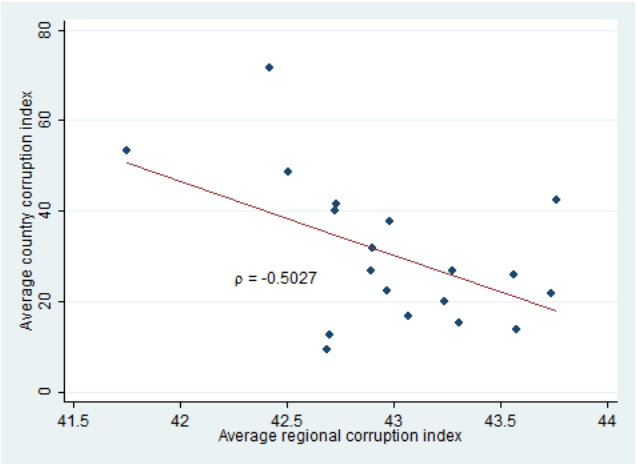

Figure 2.1: Country average versus regional average freedom from corruption index, 1995-2010: Asia

To investigate my hypothesis, I partition the sample by region and run a fixed-effects panel regression. The general configuration of the estimating equation will resemble:

$$corruption_i = \beta_0 + \beta_1 regionidx_i + \Gamma X_i + \alpha_i + \epsilon_i \tag{2.15}$$

$corruption_i$ indicates the value of the 'Freedom from corruption' index for a country i, while $regionidx_i$ indicates the value of the regional corruption index calculated in the manner described above. α_i is a country-specific effect which will control for variables that move slowly over the sample period or are time-invariant such as religion, ethnolinguistic fractionalization, and colonial history. X_i is a vector of time-varying controls for

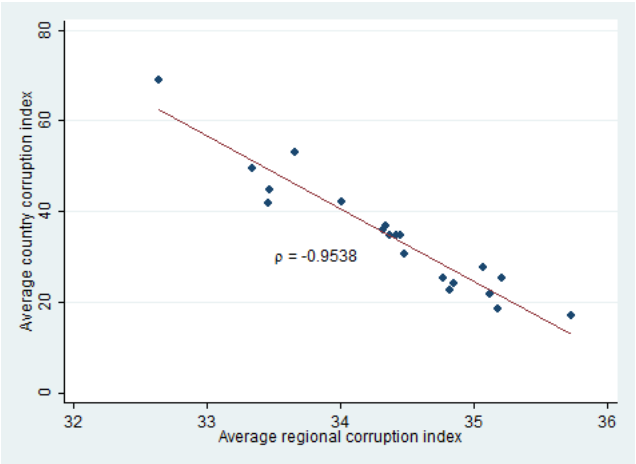

Figure 2.2: Country average versus regional average freedom from corruption index, 1995-2010: Latin America

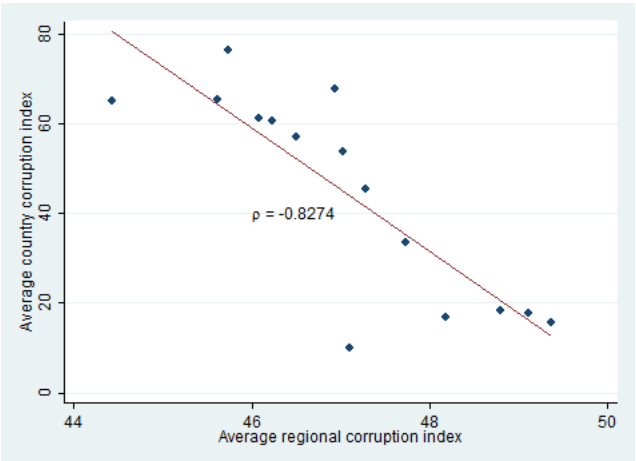

Figure 2.3: Country average versus regional average freedom from corruption index, 1995-2010: Middle East

Region	Countries/Territories	Definition
Asia	*East Asia*: China, Hong Kong, Macau, Japan, South Korea, North Korea, Mongolia. *Southeast Asia*: Cambodia, Timor-Leste, Indonesia, Laos, Malaysia, Myanmar, Papua New Guinea, Philippines, Singapore, Thailand, Vietnam. *South Asia*: Afghanistan Bangladesh, Bhutan, India, Maldives, Nepal, Pakistan, Sri Lanka.	(UN 2011)
Latin America	Belize, Costa Rica, El Salvador, Guatemala, Honduras, Mexico, Nicaragua, Panama, Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Guyana, Paraguay, Peru, Suriname, Uruguay, Venezuela.	(UN 2011)
Middle East	Bahrain, Cyprus, Egypt, Iran, Iraq, Israel, Jordan, Kuwait, Lebanon, Omar, Qatar, Saudi Arabia, Syria, United Arab Emirates, Yemen.	Traditional definition

Table 2.2: Countries/territories and corresponding geographical regions

country i. As the model indicates that consumer prices and wages are correlated with the level of domestic corruption (eq. (2.12)), I include the log of the consumer price index as a control, and proxy for wages by including the log of GDP. Data on these variables are from the *World Development Indicators*. GDP is recorded in year 2000 U.S. dollars while the consumer price index in each country is normalized to equal 100 in the year 2005. Also included in each regression are time fixed effects. Because the panel is relatively short, with at most 16 time observations for each country (corruption data begins in 1995 and national accounts data ends in 2010), I employ three-year time dummy variables in order to have a matrix of moment conditions of full rank. Given that the corruption perception index is a composite of surveys administered over multiple years (the Heritage Foundation Freedom from Corruption Index in 2010, for instance, uses surveys conducted in 2007 by the Asian Development Bank and the Political and Economic Risk Consultancy), using multiple-year time fixed effects may in fact be better than single-year time effects in capturing possible shifts in perception that might occur in the data.

To account for endogeneity due to the simultaneous determination of corruption, income and consumer prices, I instrument for GDP and CPI using data on rainfall taken from the Global Historical Climatology Network version 2 dataset, where levels of rainfall are aggregated to the annual frequency by taking average rainfall levels from weather stations within a country, summing across months, and logged. I supplement the possible instrument set with global commodity prices, including aggregate food and energy price indices, from the World Bank *GEM Commodities* Database. As commodity prices are determined by global supply and demand forces, they are exogenous from the perspective of any single country. Furthermore, I restrict attention to aggregate commodity price indices (such as food, energy and mining) rather than price series for individual commodities because a country which may have some market power by virtue of being a large supplier or demander of a particular commodity is more likely to take these aggregate price indices, which are comprised of multiple commodities, as exogenous. Commodity prices are recorded in the *GEM Commodities* database in constant year 2000 U.S. dollars at monthly frequency starting in January 1960. I aggregate to

the annual frequency by normalizing the initial monthly price to unity, averaging across months, and taking logs. For the sake of parsimony, rather than including all commodity prices as instruments along with rainfall, I choose a subset of series for each region such that the instruments appear to be uncorrelated with the error term and correlated with the endogenous regressors. In each table, I report formal test statistics for each of these assumptions.

The Reflection Problem

Another form of endogeneity that can emerge in the estimation of models in the form of equation (2.15) is known in the peer-effects literature as the reflection problem, the landmark discussion of which was given by Manski. The reflection problem "arises when a researcher observing the distribution of behaviour in a population tries to infer whether the average behaviour in some group influences the behaviour of the individuals that comprise the group." (Manski 1993) However, because the sign, rather than the magnitude, of the contagion effect is the true object of interest in this paper, there are several reasons to suggest that the reflection problem is less damaging for this investigation than for papers seeking to precisely quantify neighborhood effects. For example, consider the simplest case where two outcome variables are simultaneously determined:

$$Y_1 = \beta_1 X_1 + \rho_{12} Y_2 + \epsilon_1 \tag{2.16}$$

$$Y_2 = \beta_2 X_2 + \rho_{21} Y_1 + \epsilon_2 \tag{2.17}$$

Franzese and Hays show that OLS estimation of ρ is subject to simultaneity bias (Franzese and Hays 2007):

$$\widehat{\rho}_{12} = \rho_{12} + \frac{\rho_{21}\text{Var}(\epsilon_1)(1 - \rho_{21}\rho_{12})}{\beta_2^2 + \rho_{21}^2\text{Var}(\epsilon_1) + \text{Var}(\epsilon_2)} \tag{2.18}$$

Assuming that $\rho_{12}\rho_{21} < 1$ (cross-country influence is less than 1:1, which is empirically likely given the difficulties faced by countries in adjusting their levels of government corruption), OLS estimates of the contagion effect from country i to country j will have bias of the same sign as the contagion effect from country j to country i. If both contagion effects are of the same sign, OLS estimates will be inflated in the direction

of that sign. Positive inflation is unlikely in the results I show below, which have a negative estimate for the coefficient on the regional variable–the presence of positive bias would suggest that the true effect is even more negative. Negative inflation of the case described above would occur if the true contagion effect was indeed negative, a precondition that would reinforce the conclusions of the model of Section 2.3. If, on the other hand, feedback is dampening, and ρ_{12} and ρ_{21} are of the opposite sign, OLS estimates will be attenuated, biasing results against statistical significance. In the results I present below, the contagion coefficient is robustly significant at standard levels of statistical significance; if attenuation bias is present, it suggests the true effect is even more important than indicated.

Furthermore, one source of endogeneity that comprises the reflection problem, endogenous selection into peer groups, is absent from the context of this paper as countries are unable to choose their geographic partners. To deal with the other sources of endogeneity relating to simultaneity bias, Manski points to dynamic methods as a possible means of identifying social effects. If neighborhood effects operate on individual behavior with a lag while non-social forces act contemporaneously, the precise identification of social effects is possible. For the purposes of this paper, this entails regarding τ^* in Section 2.3 as predetermined. For the econometric results that follow, it must be emphasized that the key identifying restriction is that the true mechanism of transmission of corruption from the regional to the country level follows a recursive structure where lagged regional corruption affects contemporaneous domestic behavior. However, as the discussion above suggests, the presence of some endogeneity in the data does not entirely discredit the case being argued.

Given these considerations, the resulting estimated equation is:

$$corruption_{it} = \beta_0 + \beta_1 regionidx_{it-1} + \Gamma X_{it} + \alpha_i + \epsilon_{it} \tag{2.19}$$

2.4.2 Results

In Table 2.3, I report results for the estimation of equation (2.19) for Asia. The first column describes results when the index of regional corruption uses trade as percentage

of GDP as a measure of openness, the second uses exports as percentage of GDP, and the third employs the import-weighted regional index. Higher income levels and lower consumer prices are significantly correlated with lower corruption in Asia at $\alpha = 0.001$ and $\alpha = 0.10$, respectively. The coefficient on the regional index of freedom from corruption is significantly negative at the 1% level regardless of weighting method. An improvement in corruption amongst a country's neighbors leads to a deterioration in corruption domestically, as predicted by the model in Section 2.3.

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	20.019***	24.775***	16.095***
	(4.971)	(5.768)	(4.727)
$\log CPI_t$	-11.665+	-12.019+	-11.383+
	(6.639)	(6.722)	(6.545)
$regionidx_{t-1}$	-0.537**	-0.565**	-0.512**
	(0.175)	(0.187)	(0.168)
N	184	184	184
Clusters	15	15	15
Overall p	0.010	0.010	0.010
Weak insts. F	43.445	47.905	37.971
Hansen p	0.862	0.889	0.834

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported. Instruments are current to 2nd lag of log rainfall level and log base metal price index.

Table 2.3: Baseline results for East and South Asian countries

Table 2.4 shows analogous results for Latin American countries. Across all three specifications, the coefficient on income is negative. It is significant at the 10% level when controlling for the export weighted index and at the 5% level for the import-share weighted index, showing that in Latin America, richer countries may tend to suffer from more corruption. Moderating the effect of GDP on corruption, the estimated consumer price level effect is significant and positive across all three specifications. Even after controlling for country and time fixed effects, as well as income and CPI, the coefficient on regional corruption is significant (at $\alpha \leq 0.05$). As predicted by the model, domestic corruption varies negatively with levels of corruption amongst regional peers.

Results for the Middle East are shown in Table 2.5. For Middle Eastern countries, although the estimated coefficient is not statistically signficant, higher log GDP is correlated with greater corruption. This may reflect the fact that many Middle Eastern

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	-44.047+	-36.657	-49.303*
	(22.514)	(24.452)	(21.640)
$\log CPI_t$	22.996***	24.914***	20.878***
	(4.904)	(5.675)	(4.368)
$regionidx_{t-1}$	-2.275**	-1.753*	-2.689**
	(0.831)	(0.682)	(0.922)
N	189	189	189
Clusters	17	17	17
Overall p	0.000	0.000	0.000
Weak insts. F	12.801	15.857	10.920
Hansen p	0.372	0.454	0.302

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported.
Instruments are current to 2nd lag of log rainfall, and current and lagged log base metal prices.

Table 2.4: Baseline results for Latin American countries

countries have state-run oil- and gas-producing industries, and higher production may increase the capacity of the government to engage in graft. Higher prices are signficantly correlated with greater corruption when controlling for the trade- and export-weighted indices of regional corruption. The coefficient on the regional index of corruption is negative and significant at $\alpha \leq 0.05$ across all specifications. Higher corruption abroad leads to lower corruption domestically (and vice versa) in the Middle East as well as in Asia and Latin America. Although the effect of the control variables differs across the three regions of the world, the corruption spillover effect described in Section 2.3 emerges as robust across different methods of calculation of regional corruption, as well as across geography.[4]

2.4.3 Robustness Checks

Other time fixed effects

To assess whether the results above are robust to the specification of time fixed effects, I redo the analysis controlling for two- and five-year rather than three-year time dummy variables. The results for Asia are in Tables 2.6 and 2.7. The estimated coefficients on income and consumer prices on corruption are sensitive to the set of time effects included. For two-year time fixed effects, higher GDP is significantly associated with

[4]Results for Sub-Saharan Africa are presented in Appendix E.

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	-23.712	-24.438	-21.799
	(43.541)	(44.736)	(42.298)
$\log CPI_t$	-36.754*	-38.859*	-29.503
	(18.087)	(16.989)	(19.976)
$regionidx_{t-1}$	-0.313**	-0.295*	-0.316**
	(0.117)	(0.115)	(0.110)
N	125	125	125
Clusters	10	10	10
Overall p	0.022	0.021	0.022
Weak insts. F	48.644	43.582	46.388
Hansen p	0.572	0.590	0.504

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported. Instruments are current to 2nd lag of rain, current to 1st lag of log food prices, and current log base metal prices.

Table 2.5: Baseline results for Middle Eastern countries

greater freedom from corruption while higher consumer prices lead to a worsening of corruption. For five-year time effects, the coefficients on GDP are statistically insignificant while the coefficients on CPI reverse sign and are significant in two out of three specifications. Despite these changes, the estimated coefficient on the regional index of corruption remains negative and statistically significant. This suggests that the primary finding in the previous section regarding negative corruption spillovers is robust to alternative methods of deseasonalization, although the same cannot be said for inferences about the effect of income and prices on domestic corruption.

Results for Latin America are in Tables 2.8 and 2.9, while results for the Middle East are in Tables 2.10 and 2.11, respectively. Both tables indicate a negative effect of regional corruption on domestic corruption, supporting the hypothesis of negative spillover of corruption. The estimated coefficient on the corruption index is statistically significant in both regions and for both specifications of the time effect. That an improvement in foreign corruption can increase incentives for domestic governments to engage in graft suggests that strong enforcement mechanisms are necessary when addressing corruption in an internationally cooperative regime. Lacking such mechanisms, cooperative efforts to reduce corruption may achieve little success.

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	54.991**	57.082**	50.864**
	(18.861)	(20.226)	(18.241)
$\log CPI_t$	-13.384*	-13.359+	-12.443*
	(6.458)	(6.884)	(6.196)
$regionidx_{t-1}$	-0.491***	-0.464***	-0.508***
	(0.120)	(0.123)	(0.124)
N	245	245	245
Clusters	18	18	18
Overall p	0.010	0.019	0.009
Weak insts. F	11.163	16.083	9.951
Hansen p	0.205	0.255	0.165

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported.
Instruments are log base metal prices, current to 1st lag of log energy prices, and current to 2nd lag of log food prices.

Table 2.6: Results for East and South Asian countries with 2-year time fixed effects

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	-8.350	-5.008	-11.004
	(11.527)	(11.747)	(11.522)
$\log CPI_t$	10.946+	8.945	12.531*
	(5.691)	(5.956)	(5.571)
$regionidx_{t-1}$	-0.252*	-0.277*	-0.226+
	(0.120)	(0.124)	(0.115)
N	186	186	186
Clusters	15	15	15
Overall p	0.007	0.004	0.011
Weak insts. F	10.658	14.110	7.920
Hansen p	0.536	0.567	0.503

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported.
Instruments are % change in rainfall, current through 2nd lag of base metal prices, and log food prices.

Table 2.7: Results for East and South Asian countries with 5-year time fixed effects

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	-76.321***	-67.625***	-83.695***
	(17.080)	(17.220)	(15.670)
$\log CPI_t$	46.778***	45.076***	46.625***
	(9.983)	(10.338)	(9.101)
$regionidx_{t-1}$	-1.627**	-1.166*	-2.113***
	(0.573)	(0.489)	(0.616)
N	175	175	175
Clusters	17	17	17
Overall p	0.002	0.006	0.000
Weak insts. F	7.556	5.538	6.853
Hansen p	0.343	0.420	0.272

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported.
Instruments are current to 2nd lags of log rainfall, log food prices, and % changes in base metal prices.

Table 2.8: Results for Latin American countries with 2-year time fixed effects

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	-0.925	-5.395	1.546
	(12.730)	(11.582)	(13.853)
$\log CPI_t$	14.635**	14.449***	15.234**
	(4.450)	(4.238)	(4.750)
$regionidx_{t-1}$	-1.573**	-1.452**	-1.566**
	(0.496)	(0.443)	(0.521)
N	189	189	189
Clusters	17	17	17
Overall p	0.002	0.001	0.002
Weak insts. F	11.383	8.387	16.018
Hansen p	0.177	0.171	0.185

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported.
Instruments are current to 2nd lag of log rainfall, log food prices and log base metal prices.

Table 2.9: Results for Latin American countries with 5-year time fixed effects

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	-19.809	-14.337	-29.777*
	(12.113)	(11.353)	(12.777)
$\log CPI_t$	50.666**	60.004***	35.066*
	(15.802)	(16.410)	(13.833)
$regionidx_{t-1}$	-0.655***	-0.456***	-0.870***
	(0.155)	(0.119)	(0.163)
N	125	125	125
Clusters	10	10	10
Overall p	0.010	0.005	0.006
Weak insts. F	34.682	12.018	31.498
Hansen p	0.308	0.309	0.274

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported. Instruments are current to 2nd lags of rain, log base metal prices, and log food prices.

Table 2.10: Results for Middle Eastern countries with 2-year time fixed effects

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	-51.930	-66.143+	-45.202+
	(33.209)	(38.814)	(26.807)
$\log CPI_t$	9.862	11.614	8.046
	(18.032)	(20.712)	(14.997)
$regionidx_{t-1}$	-1.296***	-1.832***	-0.749***
	(0.213)	(0.285)	(0.204)
N	125	125	125
Clusters	10	10	10
Overall p	0.000	0.000	0.006
Weak insts. F	23.448	21.676	13.733
Hansen p	0.206	0.224	0.310

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported. Instruments are current to 2nd lag of rain and log base metal prices, and current to 1st lag of log food prices.

Table 2.11: Results for Middle Eastern countries with 5-year time fixed effects

2.5 Conclusions

Corruption has been linked to adverse consequences in economic development and the allocation of resources in many countries throughout the world. Economists who have examined how corruption is spread across borders describe a contagion effect whereby domestic corruption takes on the characteristics of corruption in geographically proximate countries. This view generally leads to optimistic conclusions about the future of corruption: as the world becomes increasingly economically and culturally integrated, countries will have greater incentives to reduce corrupt behavior in their own business and government practices, spurring similar change amongst their regional peers and collectively blazing a trail toward some corruption-free steady state.

In this paper, I build a simple model and show evidence of a corruption contagion effect that works in the opposite direction: an improvement in corruption within a region can lead to a worsening of corruption within a single country located in that region. I have shown that this spillover effect is robust across countries in Asia, Latin America, and the Middle East, even while the three regions may behave differently in response to more traditional economic determinants of corruption. My results indicate that without effective enforcement mechanisms, efforts to address corruption through intergovernmental cooperative regimes (such as the 1999 Anti-Corruption Initiative for Asia-Pacific created by the Asian Development Bank and the OECD, and the 2005 UN Convention against Corruption), may be less successful than anticipated. Corruption may be an enduring problem for socioeconomic development as nations continue to advance on the path of global integration.

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Appendix D

Corruption substitutability in general equilibrium

In this appendix I build a simple, general-equilibrium model of competitive corruption between countries, in the spirit of the classic Cournot model of industrial organization.

Consider a region of the world with n countries, each ruled by a rent-seeking government. Each government $i \in N = \{1, \ldots, n\}$ maximizes its own utility by engaging in graft, extracting rent from a representative citizen at the rate $\tau_i \in [0, 1]$. Although not essential to the analysis, I assume it is costly for the government to collect these rents. The continuous and twice-differentiable (C^2) cost function is given by $c_i(\tau_i)$ where $c_i'(\tau_i) > 0$, $c_i''(\tau_i) \geq 0$, and $c_i(0) = 0$. The representative citizen in each country earns an income of $m_i(T)$, where $m(\cdot)$ is C^2 and T represents corruption in the region as a whole.

There are many ways domestic income might be related to aggregate regional corruption. If the country is an open economy and the citizen consumes traded goods, higher regional corruption can increase the effective price of traded goods as consumers are forced to pay a corruption premium. The resulting increase in disutility then causes the citizen to substitute from consumption into leisure and earn lower income. I capture this labor-leisure tradeoff channel in reduced form by imposing the condition that income is decreasing in regional corruption or, equivalently, $m'(T) < 0$. This relationship might also arise if citizens' income contains an exogenous foreign component such as

aid, remittances, or investment that is contingent on regional corruption. For instance, foreign investors who have an imperfect capability to monitor idiosyncratic corruption may choose to give more to all countries located in regions that have lower corruption as a whole. I will additionally assume that $m''(T) < 0$; this implies a diminishing marginal impact of increased corruption on income.

A simple algebraic example of a small open economy can further highlight how this relationship between income and corruption might arise. Assume home income m is a composite of domestically produced goods and foreign imports, i.e. $m = A_1^{\rho} A_2^{1-\rho}$, where A_1 are domestically produced intermediates, A_2 are imported intermediate goods, and $0 < \rho < 1$ is the share of domestic intermediates in the income bundle. As the home economy is small in comparison to the rest of the region, we can abstract from its contribution to the rest of the world's income and assume that the latter is exogenously given as m^*, which is completely produced abroad. m^* is allocated to foreign consumption, government graft, and exports to the home economy as follows: $m^* = C^* + G^* + A_2$. Rents from foreign corruption are proportional to foreign income: $G^* = \tau^* m^*$. Using these equations, we can rewrite home income as $m = A_1^{\rho} \left((1-\tau^*)m^* - C^*\right))^{1-\rho}$. The first derivative of m with respect to τ^* is: $\dfrac{\partial m}{\partial \tau^*} = \dfrac{-A_1^{\rho}(1-\rho)m^*}{((1-\tau^*)m^* - C^*)^{\rho}} < 0$. The second derivative is $\dfrac{\partial^2 m}{\partial \tau^{*2}} = \dfrac{-A_1^{\rho}\rho(1-\rho)m^{*2}}{((1-\tau^*)m^* - C^*)^{1+\rho}} < 0$, which satisfies the conditions above. However, as it will not be necessary to specify a specific functional form for domestic income in order to derive the main prediction of the model, the remainder of the equations in this section refer to the more general set-up of the model.

Government i's utility function can be written as

$$\pi_i(\tau_i, \tau_{-i}) = \tau_i m_i(\tau_i + T_{-i}) - c_i(\tau_i) \tag{D.1}$$

where $T_{-i} = \sum_{j \in N \setminus \{i\}} \tau_j$ is a measure of corruption in the rest of the region. Given the assumptions on the income and cost functions above, there exists a pure-strategy Nash equilibrium in this game.

D.0.1 Preliminary Analysis

The first-order condition for government i's utility maximization problem is:

$$\frac{\partial \pi_i}{\partial \tau_i} = m_i(\tau_i + T_{-i}) + \tau_i m_i'(\tau_i + T_{-i}) - c_i'(\tau_i) = 0 \tag{D.2}$$

Taking derivatives again, the second order condition

$$\frac{\partial^2 \pi_i}{\partial \tau_i^2} = 2m_i'(\tau_i + T_{-i}) + \tau_i m_i''(\tau_i + T_{-i}) - c_i''(\tau_i) < 0 \tag{D.3}$$

indicates global concavity of the government's utility function. Thus, the first-order condition characterizes the best response correspondence for government i:

$$\beta_i(T_{-i}) = \{\tau_i \in [0,1] : m_i(\tau_i + T_{-i}) + \tau_i m_i'(\tau_i + T_{-i}) - c_i'(\tau_i) = 0\} \tag{D.4}$$

The best response correspondence ties domestic corruption to regional corruption as well as domestic income.

D.0.2 Strategic substitutability

The key result of the model pertains to the interaction between corruption within a country and the culture of corruption amongst its peers in its geographical region. Total differentiation of the first-order condition yields:

$$(2m_i'(\tau_i+T_{-i})+\tau_i m_i''(\tau_i+T_{-i})-c_i''(\tau_i))d\tau_i+(m_i'(\tau_i+T_{-i})+\tau_i m_i''(\tau_i+T_{-i}))dT_{-i} = 0 \tag{D.5}$$

or equivalently,

$$\frac{d\tau_i}{dT_{-i}} = \frac{m_i'(\tau_i + T_{-i}) + \tau_i m_i''(\tau_i + T_{-i})}{-(2m_i'(\tau_i + T_{-i}) + \tau_i m_i''(\tau_i + T_{-i}) - c_i''(\tau_i))} < 0 \tag{D.6}$$

which shows that this model exhibits the property of strategic substitutability often assumed in industrial organization models of Cournot competition, with the assumption of $m''(T) < 0$ serving as the analogue of the strategic substitutes assumption in the industrial organization and international trade literature.

Appendix E

Results for Sub-Saharan Africa

In the following appendix, I replicate the baseline analysis of Section 2.4 for the countries of Sub-Saharan Africa, which encompasses all of Africa except for the Northern African countries of Algeria, Egypt, Libya, Morocco, Sudan, and Tunisia. Due to data limitations, not all Sub-Saharan African countries appear in the sample. The number of countries appearing in the estimation varies with the instrument set. Table E.1 presents summary statistics for the freedom from corruption index over the sample period for all available Sub-Saharan African countries. A scatterplot of a country's average freedom from corruption index over the sample period versus its corresponding average trade-weighted index of regional freedom from corruption appears in Figure E.1. As for Asia, Latin America and the Middle East, there is a negative correlation between a country's average corruption index and the average for the corresponding index of its regional peers' corruption.

Region	Obs	Mean	Std. Dev.	Min	Max	Corr. with first lag
S.S. Africa	587	27.43271	13.11696	7	70	0.9183

Table E.1: Summary statistics for Sub-Saharan African countries' freedom from corruption indices

For the trade- and import-weighted specification, the current through first lag of the percentage change in rainfall, the current through third lag of log base metal prices, and the percentage change in energy prices were used as instruments. For the regression that included the export-weighted index, the current to first lag of percentage change

in rainfall levels, log base metal prices, and log energy prices were used as instruments for GDP and CPI. Qualitative results regarding a country's response to regional corruption are similar, but instruments become much weaker, as indicated by low values of the Kleibergen-Paap Wald F statistic. Thus, the results in this appendix are merely suggestive of the effect described in the body of this paper and are presented here for reference only. For Sub-Saharan Africa, higher income is significantly correlated with greater freedom from corruption. Higher consumer prices are associated with greater corruption, but the effect is not statistically significant. A decrease in regional corruption leads to an increase in domestic corruption as the coefficient on the regional index is negative and statistically significant at the 10% level in the trade weighted specification and at the 5% level in the other specifications.

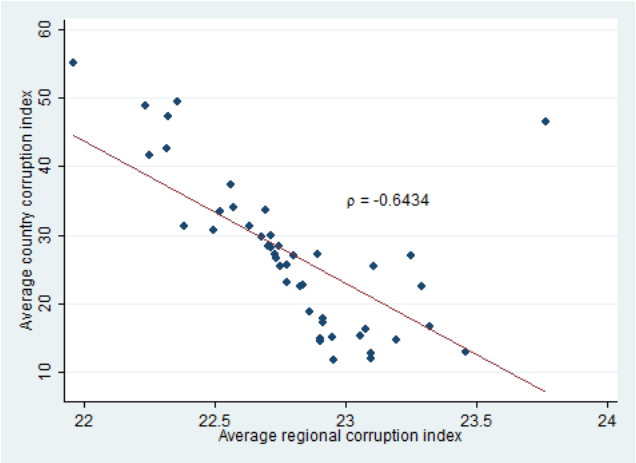

Figure E.1: Country average versus regional average freedom from corruption index, 1995-2010: Africa

Results are similar when controlling for two- and five-year time fixed effects rather than three-year time fixed effects. Tables E.3 and E.4 show that higher income is significantly correlated with lower corruption at the 5% level in all specifications. As with the baseline results, the coefficient on the regional index is negative and statistically significant. The same caveat regarding instrument strength applies to these results.

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	25.277*	30.123**	26.277*
	(11.913)	(11.196)	(12.014)
$\log CPI_t$	-9.046	-2.432	-9.730
	(5.945)	(5.127)	(6.183)
$regionidx_{t-1}$	-0.469+	-0.632*	-0.514*
	(0.246)	(0.259)	(0.258)
N	290	314	290
Clusters	31	31	31
Overall p	0.072	0.015	0.047
Weak insts. F	3.615	2.536	3.459
Hansen p	0.884	0.726	0.835

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported.

Table E.2: Baseline results for Sub-Saharan African countries

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	62.295***	63.691***	61.021***
	(12.732)	(13.780)	(12.116)
$\log CPI_t$	-6.686	-3.061	-9.030
	(7.608)	(6.788)	(8.102)
$regionidx_{t-1}$	-0.372*	-0.516*	-0.286*
	(0.178)	(0.257)	(0.132)
N	314	314	314
Clusters	31	31	31
Overall p	0.000	0.001	0.000
Weak insts. F	4.760	5.033	4.422
Hansen p	0.579	0.631	0.566

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported.
Instruments are current to 1st lag of percentage changes in rainfall, base metal prices, and food prices.

Table E.3: Results for Sub-Saharan African countries with 2-year time fixed effects

	Trade-weighted	Export-weighted	Import-weighted
$\log gdp_t$	29.063**	32.608**	25.009**
	(10.307)	(12.150)	(9.341)
$\log CPI_t$	-6.645	-6.272	-7.426+
	(4.326)	(4.388)	(4.314)
$regionidx_{t-1}$	-0.641*	-0.560+	-0.615*
	(0.300)	(0.335)	(0.263)
N	349	349	349
Clusters	33	33	33
Overall p	0.056	0.083	0.044
Weak insts. F	1.631	1.513	1.825
Hansen p	0.195	0.143	0.232

+ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. Clustered standard errors reported.
Instruments are current to 1st lag of log rainfall and log base metal prices, and current log food prices.

Table E.4: Results for Sub-Saharan African countries with 5-year time fixed effects

Chapter 3

Democracy and Dissymmetry in Interventionist Trade Policies

3.1 Introduction

Are less democratic governments more apt to intervene in the prices of imported goods than exported goods? An earlier chapter of this dissertation investigated the effect of global commodity price fluctuations on the likelihood of political conflict in lower-income countries, which are consistently ranked below richer countries on indices of political freedom. An increase in a net export-weighted agricultural commodity price index was associated with less likelihood of civil conflict, while an increase in the uncertainty of this index was associated with greater likelihood of conflict (see Table 3.1). In comparison, fluctuations in the level and uncertainty of an analogous net import-weighted index were not found to be statistically significant predictors of civil conflict events (Table 3.2). The asymmetric effects of global commodity import and export prices on domestic political conditions suggest that in this sample of countries, domestic prices of imported commodities are less tightly linked to global prices than domestic prices of export commodities.

In this paper I offer an explanation for why this might be the case. I focus on a government's choice between two alternative interventionist trade policies: import tariffs and export subsidies. If governments have incentives to exploit their political

	LPM-RE	LPM-RE	IV-LPM-FE	IV-LPM-FE	LOGIT-FE	PROBIT-RE
ΔPI_t	-0.036*	-0.043*	-0.060**	-0.041+	-0.404+	-0.217+
	(0.018)	(0.019)	(0.021)	(0.022)	(0.221)	(0.128)
$\Delta \text{Vol}_1\text{PI}_t$	0.148+	0.160+	0.177+	0.186*	1.980+	1.035+
	(0.085)	(0.089)	(0.107)	(0.092)	(0.994)	(0.560)
Δrain_t		0.002				
		(0.001)				
$\Delta \text{reserves}_t$			-0.029			
			(0.037)			
Δincome_t				-0.194	-2.380	-1.308
				(3.328)	(1.902)	(0.963)
Constant	0.194***	0.198***				-1.645***
	(0.029)	(0.031)				(0.319)
$\ln \sigma_u^2$						1.416***
						(0.343)
N	13950	12621	9399	9131	6576	10385
Clusters	75	74	69	73	48	75
Overall p	0.077	0.074	0.039	0.133	0.086	0.096
Over-i.d. test p	0.421	0.481	0.258	0.224		
Endog. test p			0.367	0.449		

$+\ p < 0.10$, $*\ p < 0.05$, $**\ p < 0.01$, $***\ p < 0.001$
Cluster-robust standard errors reported in parentheses.

Table 3.1: Estimation using price variance to measure price uncertainty, net export–weighted indices

power to extract rents from citizens, they can achieve this end by taxing imports rather than subsidizing exports. However, if citizens are able to discipline their governments through elections, the extent of this rent-seeking behavior can be constrained. The greater the extent of electoral accountability in a country, the smaller its tariff rents will be. This implies that more democratic countries will be less willing to tax imports than authoritarian countries. Authoritarian governments in the regressions in Tables 3.1 and 3.2, using tariffs to drive a wedge between the domestic price of imported commodities and the global price of these commodities, may explain the lack of significance in the relationship between global commodity import prices and intrastate civil conflict, even as a relationship exists on the export side of the economy.

I present a model that captures the economic intuition above. One feature of the model is that it distinguishes between the level of electoral accountability of a government to its citizens from the level of bargaining power that a government has over a citizen lobby. Until now, this distinction has not been incorporated into a model that determines levels of both tariffs and subsidies. While some authors argue that citizen bargaining power reflects the openness of a political system, I show that the level of

	IV-LPM-FE	LPM-FE	LOGIT-RE	LOGIT-FE	PROBIT-RE
$\Delta\text{PI}_{imp,t}$	0.004	0.010	0.116	0.116	0.076
	(0.030)	(0.027)	(0.317)	(0.315)	(0.171)
$\Delta\text{VarPI}_{imp,t}$	0.023	0.014	0.193	0.188	0.095
	(0.061)	(0.054)	(0.632)	(0.627)	(0.344)
Δincome_t	-0.697	-0.299	-2.506	-2.455	-1.330
	(2.322)	(0.192)	(1.814)	(1.781)	(0.910)
Constant		0.224***	-3.028***		-1.660***
		(0.000)	(0.532)		(0.304)
$\ln \sigma_u^2$			2.717***		1.359***
			(0.311)		(0.327)
N	10226	11829	11829	7512	11829
Clusters	84	87	87	56	87
Overall p	0.667	0.231	0.297	0.304	0.277
Over-i.d. test p	0.209				
Endog. test p	0.969				

$+\ p < 0.10$, $*\ p < 0.05$, $**\ p < 0.01$, $***\ p < 0.001$

Cluster-robust standard errors reported in parentheses.

Table 3.2: Effect of net import-weighted price level, volatility, and income on conflict incidence

democracy and the level of citizen bargaining power can have opposite effects on the tariff rate when tariff revenues fund both government rents and export subsidies. After establishing the main conclusion that import tax expenditure is greater than export subsidy expenditure in more authoritarian countries, I present some empirical evidence in support of this conclusion, showing that the difference between import tariff and export subsidy expenditure is correlated with standard indices of political rights and democracy. These findings give insight into some of the challenges of establishing free trade amongst countries with different attitudes toward democratic institutions. Given the persistence of political institutions, these results also offer an alternative explanation for why multilateral trade organizations such as the GATT/WTO, whose membership consists of countries with widely different political regimes, tend to be more strict on export subsidies than on import tariffs.

The remainder of this paper is organized as follows. The next section presents a review of a sample of literature related to the question at hand. The theoretical model is presented in Section 3.3. Empirical findings presented in Section 3.4 lend support to

the main conclusions of the model. Section 3.5 offers concluding remarks.

3.2 Literature Review

Research into the political economy of trade is extensive, and I make no attempt in this paper to provide a complete survey of the field. A broader survey of the existing literature is given by (Rodrik 1995), among others. Instead, in this section I discuss a small sample of work which is related to my investigation, and discuss the differences in these approaches with respect to my own.

This paper is related to the literature which examines how trade policy can be determined in a political equilibrium. Findlay and Wellisz develop a simple model of trade where tariffs are endogenously determined through a political process (Findlay and Wellisz 1982). In their model, landed interests who favor a tariff on food compete with a manufacturing sector which favors free trade over the political determination of the tariff level. The conflict of interest between the two sectors is due to the reallocation of labor costs that would emerge as the country opens to trade: the rental cost of capital in the manufacturing sector, which has comparative advantage, would decrease under trade while the labor cost in the farming sector would increase. The primary finding of their model is that tariffs that arise endogenously impose a larger welfare cost than exogenously given tariffs, which had previously been assumed in much of the literature.

Hillman considers how governments motivated by the pursuit of political support implement protectionist tariff policies for industries in decline (Hillman 1982). The author adapts the Stigler–Peltzman regulatory model to a Ricardo–Viner trade setting, finding that when authorities are motivated by political self-interest, they will not fully offset welfare losses faced by an industry contracting due to global market pressures. Depending on the political weight of the industry, the government may use tariff policy to either accelerate or retard the decline in the industry, but the industry's direction of change is predetermined by its comparative advantage in the world market.

Yang develops a model of representative democracy where candidates for political office have incomplete information regarding voters' nonpolicy preferences (Yang 1995). He shows that the framework gives rise to a probabilistic voting model where the equilib-

rium tariff is a weighted mean of voters' most preferred tariffs, and that the resulting income distribution is more equal than under free trade. While (Findlay and Wellisz 1982), (Hillman 1982), and (Yang 1995) abstract from nontariff policies, in this paper I will consider how political competition can influence the relative sizes of import tariff and export subsidy policies. Additionally, while the model in my paper will also allow for varying levels of democracy in a country, the previous papers assume the existence of a consolidated democracy or the government's need for citizen support.

Maggi and Rodriguez-Clare develop a small-country model to show how trade agreements which enforce free trade may yield higher welfare than the political equilibrium characterized by lobbying between interest groups and governments (Maggi and Rodriguez-Clare 1998). In their paper, a manufacturing sector will lobby with the government for protectionist trade policy while the government maximizes a weighted average of total welfare and lobbying contributions. The authors show that governments may wish to commit to free trade in order to insulate themselves from domestic political pressures, but the stronger the government, the less likely it will be to engage in a free trade agreement, due to a greater ability to extract higher rents from the political process. While Maggi and Rodriguez-Clare's model features a single industrial sector lobbying for only one kind of protectionist trade policy, my paper will allow both import tariffs and export subsidies to exist simultaneously.

In a later paper, Maggi and Rodriguez-Clare describe a novel model that pins down optimal levels of import tariffs and export subsidies in trade agreements (Maggi and Rodriguez-Clare 2005). They augment a standard trade model by allowing incomplete trade agreements where agreements can specify tariff and subsidy ceilings, rather than exact levels, and by allowing for both ex-ante bargaining over trade policy ceilings and ex-post bargaining over levels of protectionism subject to the constraint of the ceiling. A key parameter in their paper reflects the strength of ex-ante bargaining relative to ex-post bargaining. Weaker ex-ante bargaining implies that it is preferable to set policy ceilings rather than exact policy commitments, and allows optimal levels of protectionist policies to be determined.

Potipiti directly addresses the issue of uneven implementation of trade policy in a

manner that is biased against trade (i.e. favoring import tariffs rather than export subsidies) for countries that are members of the WTO (Potipiti 2012). The author develops a small open economy model with two completely separable sectors, one engaging in the production of an import-competing good and the other producing an export good. Both sectors can lobby the government for trade protection, while the government maximizes a weighted sum of social welfare and political contributions. If capital investment in the import-competing sector is sunk and transportation costs are decreasing, there will be growth in the export sector but not the import sector. Entry in the export sector will reduce the amount of rents the government can obtain from subsidizing exports. Sunk capital in the import sector allows the rate of return on protection to increase without attracting new entrants. In Potipiti's model, if the government has sufficient bargaining power relative to the lobbies then an optimal international trade agreement will allow tariffs but prohibit subsidies.

In contrast to (Maggi and Rodriguez-Clare 1998), (Maggi and Rodriguez-Clare 2005) and (Potipiti 2012), my model will distinguish between lobbying power and electoral accountability. For example, a democratic government can have high levels of electoral accountability in that citizens are able to re-elect new administrations when previous policies fail to achieve expectations; at the same time, authorities may also be highly indebted to special interests and elite groups. However, a dictatorship where citizens are forced to acquiesce to policies enacted by the government without the ability to exercise a political role can also be dependent on the support of special interest groups. Conversely, it may also be the case that a democracy can be less susceptible to special interests as voters punish the government for favoring a privileged few over the interest of the many, while the institutions of authoritarianism may insulate a government from pressure by lobbying groups. My model explicitly takes into account the level of electoral accountability in the country and distinguishes it from the level of bargaining power and influence of special interests.

This paper is also related to the empirical literature which tries to explain the difference in the implementation of trade policies across countries. Anderson constructs and calibrates a computable general equilibrium model of trade between a representa-

tive rich and poor country to explain why poor countries protect their industrial sectors at the expense of larger agricultural sectors while rich countries favor protecting their farming industries (Anderson 1995). In Anderson's model, when agricultural prices rise, wages and the price of nontradables show a large increase in the poor economy than in the rich economy. Under the assumption of short-run factor immobility, the increase in wages attenuates the rise in factor rewards in the agricultural sector and amplifies the reduction in factor rewards in the industrial sector. Price supports thus have a smaller growth effect in the agricultural sector and a more deleterious effect in the industrial sector in rich than poor countries, which explains why they are far less prevalent than protectionist policies targeted toward the infant industrial sector. In comparison, the empirical results shown in this paper are the result of an econometric analysis rather than a calibration exercise.

Mansfield and Busch examine how various political variables relate to the incidence of nontariff barrier across countries (Mansfield and Busch 1995). Their focus is on nontariff barriers applied to imports, such as import quotas, levies, and trade restraints under the Multifibre Arrangement. Their analysis finds a relationship between macroeconomic fluctuations, particularly as captured by unemployment and exchange rate fluctuations, and levels of protectionist policy. They also find that nontariff barriers are higher in states where policymakers are most insulated from interest-group pressures. Unlike Mansfield and Busch, I study how political institutions affect the disparity between alternative trade policies. Additionally, I offer a theoretical model which lends some guidance on factors which might influence these policies, while Mansfield and Busch's analysis is loosely based on a review of the political science and political economy literature.

The choice between alternative modes of trade protectionism is addressed in a different manner by (Bown 2004). Bown analyzes trade disputes under the GATT regime from 1947 to 1994, distinguishing between episodes where protectionist tariffs were implemented legally (i.e. authorized under the GATT/WTO safeguards provisions) or illegally (without GATT notification). The author finds evidence that illegal policies are more likely when the country being targeted by the protectionist policy has import-

competing industries that are highly protected through non-tariff barriers, when the target country receives few of the protectionist-implementing country's exports, and when the protectionist-implementing country's protected sector has low pre-existing levels of trade protection. Legal protection is more likely to be employed when the protectionist-implementing country needs to shield itself from retaliation by the target country in a trade dispute. In contrast to Bown's approach which divides trade policy into legal versus illegal adjustments of GATT/WTO tariff bindings, this investigation is primarily interested in differences between tariff and subsidy volumes across countries, and the political determinants thereof.

Within the empirical literature, this paper most closely fits amongst those studies that directly analyze the impact of democracy on trade policies. Kono gives evidence that democracy leads to lower tariffs, higher core nontariff barriers, and much higher quality nontariff barriers (Kono 2006). He argues that this result is due to less transparency in the welfare effects of nontariff barriers versus tariffs. While democracy generally increases policy transparency in a country, it can also lead policymakers to substitute from more transparent trade barriers to less transparent ones. In a later paper, Kono finds empirical evidence that democratization can cause a country to adopt more liberal trade policies toward richer countries but not poorer ones (Kono 2008). His findings can be rationalized by a standard Heckscher–Ohlin–Vanek model where democracy shifts political influence from capital to labor. In this framework, labor would benefit from lower tariffs on capital-intensive goods produced by rich countries and higher tariffs on labor-intensive goods produced by poor countries. In comparison to the papers by Kono, I offer a theoretical model and place more emphasis on export subsidy policies.

The next section describes a simple model of political interaction between citizens and the government in an open economy. The model highlights how political variables, in particular the levels of electoral accountability of governments and bargaining power of citizens, can affect expenditure on export subsidies and import tariff revenues.

3.3 Model

The model in this paper features a small open economy in which prices of imports and exports are determined in the rest of the world and are treated as exogenous from the perspective of domestic agents. The economy is populated by a large number of identical voters on the continuum $[0, 1]$, as well as a policymaker. Voters and policymakers engage in a political game over fiscal and protectionist policy. Money is introduced through a cash-in-advance constraint that requires citizens to conduct transactions in foreign currency; money also appears in the utility function of the government. Adopting the small open-economy framework implies an abstraction from the strategic use of trade policy between countries, instead emphasizing the interrelationship between trade policy and domestic political institutions.[1]

The game consists of the following process of events:

1. The state of the world, which is comprised of the world prices for the import and export good ($[P_F^*,\ P_H^*]$, respectively) and the citizens' commodity endowment Y_H, is revealed.

2. The policy vector, comprised of export subsidies and income taxes (s, τ), is determined in a two-stage political game. In the first stage, citizens appoint a lobbyist to negotiate with the government over the size of the export subsidy. In the second stage, voters and the government play a political game whereby the citizens set a reservation level of utility for supporting the government, while the government simultaneously chooses a tax rate to impose on the citizens.

3. Citizens engage in trade while the government implements the announced policies.

4. Elections are held where the voters choose between the incumbent policymaker and an identical opponent.

5. Election results are respected with probability $\pi \in [0, 1]$.

[1] In a general-equilibrium model with strategic international trade policy, the welfare effects of export subsidies would depend on the nature of competition between countries. (Brander and Spencer 1985) show that a subsidy to exports raises the home country's welfare under Cournot duopoly, while (Eaton and Grossman 1986) prove that under Bertrand duopoly, a tax on exports would increase the welfare of the home country.

3.3.1 Citizen behavior

The utility of the representative citizen is:

$$U(C_F) = C_F \tag{3.1}$$

Citizens are subject to a cash-in-advance constraint that requires them to use money (denominated in foreign currency) to buy consumption goods. Foreign currency is obtained by exporting their commodity endowment to the rest of the world:

$$(P_H^* + s)Y_H = M \tag{3.2}$$

where P_H^* is the price of the export good and s is the specific subsidy, in units of foreign currency.

Citizens' export profits are subject to taxation by the government at rate $\tau \in [0, 1]$. The after-tax export profits of the citizen are used to purchase the imported consumption good:

$$P_F^* C_F = (1 - \tau)M \tag{3.3}$$

where P_F^* is the price of the import good in units of foreign currency.

Equations (3.2) and (3.3) give the citizens' budget constraint as:

$$(P_H^* + s)Y_H = \frac{P_F^*}{(1 - \tau)} C_F \tag{3.4}$$

Citizens have a limited ability to punish their governments for failing to guarantee them a minimum standard of living $\overline{\omega}$. If citizens' utility does not achieve this threshold level, they will withdraw support from the government and, conditional on the country's electoral results being respected, replace the incumbent policymaker with an identical opponent. The only incentive of citizens to change their government is to punish the incumbent. Given that the incumbent's political opponent is identical, this punishment mechanism is weakly optimal to non-punishment of the regime. Let the citizens' political

support decision be represented by the dummy variable p_I.

$$p_I = \begin{cases} 0 & \text{if } U(C_F) < \overline{\omega} \\ 1 & \text{if } U(C_F) \geq \overline{\omega} \end{cases} \tag{3.5}$$

3.3.2 Government behavior

The expected utility of the government is:

$$E[v_D] = \pi\left(r + p_I R\right) + (1 - \pi)\left(r + R\right) - sY_H \tag{3.6}$$

Election results are respected with probability $\pi \in [0, 1]$, a measure of the strength of democratic institutions in the country. The government derives utility from its revenues r as well as disutility from the subsidy to exporters, which is financed through international borrowing and must be paid back at the end of the game. If election results are respected and the citizens choose to support the incumbent, the government will receive a utility increase of R, an exogenous benefit of holding office. If election results are not respected, the government will receive this utility boost regardless of the citizens' voting behavior.

Government revenues are equal to the citizens' export revenue multiplied by the tax rate.

$$r = \tau M = \tau(P_H^* + s)Y_H \tag{3.7}$$

The structure of the game implies that the government faces a choice in elections between two options: it can judiciously choose tax policies in an attempt to remain in power, or it can abuse its position to maximize its rents, accepting that citizens will choose to expel it from office in the election. If the government chooses to not satisfy its constituents, it will impose the maximum tax rate, appropriating all export sales proceeds and behave as though voters will no longer comply. Expected utility from this

decision is:

$$\begin{aligned} E[v_D^0] &= \pi\left(r + p_I R\right) + (1-\pi)\left(r + R\right) - sY_H \\ &= \pi\left(M + (0)R\right) + (1-\pi)\left(M + R\right) - sY_H \\ &= \pi(P_H^* + s)Y_H + (1-\pi)\left((P_H^* + s)Y_H + R\right) - sY_H \\ &= P_H^* Y_H + (1-\pi)R \end{aligned} \tag{3.8}$$

On the other hand, if the government aims to acquire citizen support:

$$\begin{aligned} E[v_D^1 | p_I = 1] &= \pi\left(r + (1)R\right) + (1-\pi)\left(r + R\right) - sY_H \\ &= r + R - sY_H \end{aligned} \tag{3.9}$$

I assume that a policy commitment mechanism is in force in the economy, so that even if election results are not respected, the government will implement its announced policy vector rather than defecting and imposing the maximum tax rate.

The government will choose to acquire citizen support if:

$$\begin{aligned} r + R - sY_H &\geq P_H^* Y_H + (1-\pi)R \\ r &\geq (P_H^* + s)Y_H + (1-\pi)R - R \\ r &\geq (P_H^* + s)Y_H - \pi R \end{aligned} \tag{3.10}$$

If (3.10) does not hold, then the tariff rate does not move with political variables. [2] The interesting case occurs when the parameters of the model satisfy (3.10), which is an assumption that will be maintained for the course of this paper.

3.3.3 Equilibrium

Determination of the subsidy

To determine the subsidy rate for the export good, the citizens select a single representative to negotiate or lobby with the government. The lobbyist's objective is:

$$U_L = (P_H^* + s)^{\phi_L} \tag{3.11}$$

[2] If $\tau = 1$, the implied tariff is infinite.

which is increasing in the size of the subsidy.

In negotiations, the government's objective is decreasing in the size of the subsidy.

$$U_G = (\nu - s)^{\phi_G} \tag{3.12}$$

$\phi_L,\ \phi_G \geq 0$ are parameters which measure the relative bargaining powers of the lobbyist and the government in negotiations, respectively. The punishment utilities for not reaching an agreement are $(0, 0)$, giving incentive for both parties to reach an agreement. ν is the highest possible subsidy rate that the government will consider. In (3.2), it is assumed that the government can engage in short-term borrowing to finance the export subsidy. Borrowing occurs under the constraint that all borrowing is paid back to foreign creditors at the end of the game, which is incorporated into the government's utility function (3.6). Thus, ν can be thought of as a credit constraint parameter à la (Kiyotaki and Moore 1997); in this case, the parameter measures the total amount of foreign currency the government can borrow per unit of export good. In negotiations, the government would prefer not to borrow the maximum amount (i.e. so that the credit constraint binds), due to a desire to maintain precautionary savings for counteracting adverse macroeconomic fluctuations.

The lobbyist and the government engage in a Nash bargaining game. The Nash product $\mathcal{N}$ is:

$$\mathcal{N} = (P_H^* + s)^{\phi_L}(\nu - s)^{\phi_G} \tag{3.13}$$

The first-order condition for maximization of the expression in (3.13) is:

$$\frac{d\mathcal{N}}{ds} = \phi_L(P_H^* + s)^{\phi_L - 1}(\nu - s)^{\phi_G} - \phi_G(P_H^* + s)^{\phi_L}(\nu - s)^{\phi_G - 1} = 0 \tag{3.14}$$

Solving for s in (3.14) yields the equilibrium level of subsidies:

$$s = \frac{\phi_L \nu - \phi_G P_H^*}{\phi_L + \phi_G} \tag{3.15}$$

s is a function of the constant parameters of the model. In order to ensure that s is in fact a subsidy and not a tariff, it is assumed that $\phi_L \nu - \phi_G P_H^* \geq 0$. (3.15) shows

that, *ceteris paribus*, the greater the bargaining power of the lobbyist ϕ_L, the closer the equilibrium subsidy will be to the borrowing constraint. The greater the bargaining power of the government ϕ_G (keeping ϕ_L constant), the less it will choose to borrow.

Determination of the tax rate

While the lobbying process determines the export subsidy implemented by the government, the income tax rate and implied import tariff rate are pinned down by the electoral accountability mechanism. Because the government wishes to remain in power, it will set tax rates as high as possible so that citizens will still achieve their reservation level of utility for compliance with the government.

$$\begin{aligned} C_F &= \overline{\omega} \\ (1-\tau)\tfrac{(P_H^*+s)}{P_F^*}Y_H &= \overline{\omega} \\ \tfrac{(P_H^*+s)}{P_F^*}Y_H - \tfrac{r}{P_F^*} &= \overline{\omega} \\ &\Rightarrow \\ r &= (P_H^*+s)Y_H - P_F^*\overline{\omega} \end{aligned} \tag{3.16}$$

Citizens, on the other hand, want tax rates (and, by extension, the government's rents) to be as low as possible. They will set their reservation level of utility so that the government's power-holding constraint (3.10) holds with equality:

$$r = (P_H^* + s)Y_H - \pi R \tag{3.17}$$

(3.16) and (3.17) pin down the reservation level for citizen compliance with the government as:

$$\overline{\omega} = \frac{\pi R}{P_F^*} \tag{3.18}$$

This shows that in more democratic countries (higher π), citizens are better able to demand a higher minimum standard of living from their government.

Combining (3.7) and (3.17) gives the equilibrium tax rate:

$$\tau^* = 1 - \frac{\pi R}{(P_H^* + s)Y_H} \tag{3.19}$$

In the equilibrium of the model, the incumbent policymaker will always be re-elected, and thus will stay in office whether or not election results are respected. In the equilibrium, the government will receive the utility $v_D^* = r^* + R - sY_H = P_H^* Y_H + (1-\pi)R$. Combining (3.16) and (3.18) shows that citizens will receive utility $U = \frac{\pi R}{P_F^*}$.

3.3.4 Tariff levels and democracy

Note from (3.3) that by taxing income, the government chooses the effective price of imports faced by citizens, where the effective price is $\frac{P_F^*}{(1-\tau)}$. As P_F^* is the price of the imported good, choosing a tax rate is equivalent to choosing an import tariff. The implied ad valorem tariff rate t can be solved for as:

$$\begin{aligned} (1+t)P_F^* &= \tfrac{P_F^*}{(1-\tau)} \\ \rightarrow t &= \tfrac{\tau}{1-\tau} \end{aligned} \tag{3.20}$$

where $\frac{dt}{d\tau} = \frac{1}{(1-\tau)^2} > 0$, so an increase in the tax rate can be thought of an as increase in the tariff.

Taking the derivative of (3.19) with respect to the level of democracy π gives $\frac{d\tau^*}{d\pi} = \frac{-R}{(P_H^*+s)Y_H} < 0$. As the level of democracy increases, tariffs decrease. The model suggests that more democratic countries will have smaller levels of tariffs relative to subsidies, and that dictatorships will exploit the lack of political freedom in their countries to extract higher import tariff revenues relative to subsidy expenditures.

Democracy vs the bargaining power of citizens

The model also highlights the difference between the effect of the level of democracy in a country and the relative bargaining power between the government and the export lobby on trade policy. In the literature on the political economy of trade, these two effects are sometimes confounded in a single parameter which in reality only reflects bargaining power. For example, in (Maggi and Rodriguez-Clare 1998), the authors "do not attempt to model the factors affecting government's bargaining position, but...have in mind that such a position may be weaker in countries with more open political systems." (p. 577)

In contrast, consider the effect of an increase in citizens' relative bargaining power

on the equilibrium tariff in the present model. An increase in the export lobby's bargaining power ϕ_L will increase the size of the equilibrium export subsidy; from (3.15), $\frac{ds}{d\phi_L} = \frac{\phi_G(\nu+P_H^*)}{(\phi_L+\phi_G)^2} > 0$. In turn, an increase in the size of the export subsidy will lead to an increase in the tax rate (and by extension, import tariffs), as the government must extract more rents to finance the subsidy: from (3.19), $\frac{d\tau^*}{ds} = \frac{\pi R Y_H}{[(P_H^*+s)Y_H]^2} > 0$. Thus there is a direct relationship between the citizens' bargaining power vis–à–vis the government and the prevailing level of import tariffs. However, as discussed above, an increase in the level of democracy π will be associated with a decrease in the size of the import tariff in equilibrium: from (3.19), $\frac{d\tau^*}{d\pi} < 0$. This result emphasizes that bargaining power and political accountability can have distinct effects on interventionist trade policy.

3.4 Empirical Analysis

An implication of the model presented in the previous section is that more democratic countries will have a lower rate of expenditure on import tariffs to export subsidies. Does this prediction appear to hold true in the data? In this section I present some empirical evidence for a systematic relationship between the strength of democratic institutions and differences in the implementation of tariffs and subsidies.

3.4.1 Data

To measure the strength of democratic institutions in countries, I rely on two indices of political freedom: the 2013 Freedom in the World Index published by Freedom House and the Polity2 Score from the Polity IV dataset. The FIW index is based on a yearly survey of regional experts and scholars regarding political rights (electoral process, political pluralism, and government functioning) and civil liberties (freedom of expression and belief, associational/organizational rights, rule of law, personal autonomy and individual rights). The index, which assigns each country a score from 1 (most free) to 7 (least free) has been published annually since 1972. The Polity Score covers all independent states from 1800-2011 and is "also the most closely scrutinized data series on political issues as analysts and experts in academia, policy, and the intelligence community regularly

examine and often challenge Polity codings."[3] Each country is evaluated on the basis of executive recruitment, constraints on executive authority, and political competition and is assigned a score from -10 (hereditary monarchy) to +10 (consolidated democracy).

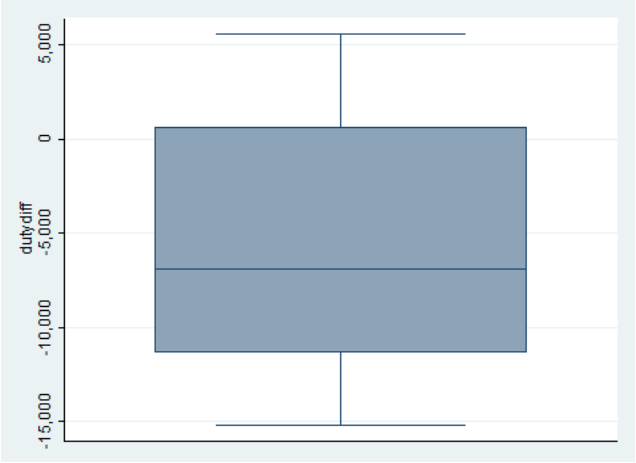

Figure 3.1: Dispersion of difference between import tariff and export subsidy expenditure.

Data on tariffs is obtained from the World Bank's *World Development Indicators*, where revenue on customs and import duties in current local currency units is available at the country level for each year from 1960–2012. I convert all values into U.S. dollar amounts by dividing by the dollar exchange rate (also available from the WDI). Data on export subsidies is taken from the WTO; I use the series on WTO members' export subsidy commitments by product and tariff line, aggregating bound values across products and converting the sum to U.S. dollar amounts. Data on export subsidy commitments is not available for each country, and I leave missing values as missing rather than assuming countries impose zero subsidies when data is unavailable. Although subsidy commitments are in actuality an upper bound which may not be the actual export subsidy that is applied, this series can nevertheless be regarded as a proxy for the true value of export subsidy expenditure in the reporting country. Data on real GDP per capita in year 2000 U.S. dollars, fertility rates, export value indices, and the trade share (sum of exports and imports) of GDP are from the World Development Indicators. Annual data on country-level export price indices are from the IMF International Financial Statistics. Summary statistics for each variable are in Table 3.3.

[3]From Polity IV project website: http://www.systemicpeace.org/polity/polity4.htm.

Variable	Obs	Mean	Std. Dev.	Min	Max
Tariffs minus subsidies	136	-5775.82	6645.23	-15206.74	5551.99
findex	6884	3.75	2.05	1.00	7.00
polity2	7104	0.65	7.48	-10.00	10.00
Log GDP per capita	9462	7.67	1.58	4.00	11.59
Export value index	5313	142.65	175.88	0.29	3951.42
Export price index	954	67.78	41.54	0.12	179.55
Military sophistication	3523	9.17	1.48	4.08	13.37
fertility	10947	4.12	1.99	0.85	9.22
Trade share of GDP	8895	73.77	47.81	0.31	460.47
Import share of GDP	8896	39.40	25.33	-17.14	221.51

Table 3.3: Summary statistics for data

3.4.2 Empirical evidence

Variables	polity2	findex	imp_duties	subsidy	dutydiff
polity2	1.000				
findex	-0.892	1.000			
imp_duties	-0.051	0.036	1.000		
subsidy	0.290	-0.427	-0.494	1.000	
dutydiff	-0.364	0.567	0.614	-0.990	1.000

Table 3.4: Cross-correlation table

Table 3.4 displays unconditional correlations between the political indices (*polity2, findex*) and import tariff (*imp_duties*) and export subsidy (*subsidy*) variables, where *difference* is the difference between revenue collected from import tariffs and a country's export subsidy commitments. There is a high correlation between the two indices of democracy; the correlation is negative due to the way each index is defined. The correlation between the democracy indices and revenue from customs and import duties indicates that more democratic countries collect smaller levels of tariffs, as predicted by the model. Examining the unconditional correlation between the democracy indices and the difference between tariff revenues and subsidy expenditures shows that more democratic countries impose smaller levels of import tariffs relative to export subsidies.

Table 3.5 shows results from an OLS panel regression of the difference between tariff and subsidy expenditure on the democracy indices, partialling off the contribution of time fixed effects and country fixed effects. Following much of the empirical literature

on the effects of democracy on trade, I treat the democracy index as an exogenous independent variable in these regressions. Csordás and Ludwig find no evidence of causality in the reverse direction: a country's openness to trade is not a determinant of its level of democracy (Csordás and Ludwig 2011). Across all columns of Table 3.5, the estimated coefficient is what the model would predict: more democratic countries spend less on tariffs than subsidies. Allowing the country effect to be correlated with the democracy index does not change the direction of this correlation, although the estimated coefficient of the Freedom House index is not significant in the fixed-effect model (REG4).

	REG1	REG2	REG3	REG4
polity2	-186.67***	-120.03*		
	(70.92)	(60.21)		
findex			1148.83*	706.61
			(673.75)	(667.97)
Observations	133	133	136	136
R-squared	0.34	0.27	0.54	0.44
p-value	0.00	0.00	0.00	0.00
Time dummies	Yes	Yes	Yes	Yes
Indiv. effect	Random	Fixed	Random	Fixed

* p<0.1, ** p<0.05, *** p<0.01. Cluster-robust standard errors reported.

Table 3.5: Panel regression with only democracy index

Although Tables 3.4 and 3.5 are indicative of some relationship between political freedom and variation in expenditure on tariffs and subsidies across countries, it would be preferable to see whether this relationship is robust to the inclusion of some country-specific time-varying controls. In the remainder of this section, I augment the baseline specification in Table 3.5 incrementally, leading to the final regression specification in Table 3.12. Equations (3.15) and (3.19) offer some guidance on possible controlling variables, indicating that import tariff and export subsidy rates are dependent on:

1. Level of income Y_H
2. Price of export good P_H^*
3. Exogenous value of holding office R
4. Credit constraint parameter ν

5. Relative bargaining power of lobbies and government ϕ_L, ϕ_G

Under the assumption that items 3 through 5 in the above list are constant over the sample period, they will be captured in the country effects. The basic OLS regression is augmented with series on each country's income per capita and export value index in Table 3.6. Controlling for these series does not change the qualitative implications regarding the relationship between political institutions and trade policies. More authoritarian governments appear more likely to employ import tariff measures than export subsidies. While export value indices are not equivalent to export price indices because they reflect changes in both the prices and composition of exports, they are nevertheless used as surrogates for price indices by many countries.[4] As a robustness check, Table 3.7 controls for the export price index series maintained by the IMF. Although there is a large drop in the number of observations available, the conditional correlations are qualitatively similar.

	REG5	REG6	REG7	REG8
polity2	-178.49***	-129.89***		
	(67.64)	(40.89)		
findex			1188.21*	1013.30***
			(662.12)	(301.98)
Log GDP per capita	11.55	11620.93***	1090.45	13084.39***
	(975.74)	(1663.20)	(843.74)	(2008.55)
Export value index	-17.01	-30.40***	-15.01	-29.75***
	(14.15)	(9.28)	(13.40)	(8.74)
N	133	133	136	136
R-squared	0.35	0.06	0.47	0.03
p-value	0.00	0.00	0.00	0.00
Time dummies	Yes	Yes	Yes	Yes
Indiv. effect	Random	Fixed	Random	Fixed

* p<0.1, ** p<0.05, *** p<0.01. Cluster-robust standard errors reported.

Table 3.6: Panel OLS augmented with income and export value index

To address possible bias in the estimate on the coefficient of the democracy index due to endogeneity between the dependent variable and the additional control variables, Tables 3.8 and 3.9 present analogous results to Tables 3.6 and 3.7, instrumenting for the log of GDP per capita and the export value index or export price index. The first

[4] http://www.imf.org/external/np/sta/tegeipi/ch2.pdf.

	REG9	REG10	REG11	REG12
polity2	-1787.30***	-1464.58***		
	(124.41)	(56.47)		
findex			3770.04***	3585.09***
			(207.80)	(367.01)
Log GDP per capita	427.40	3236.08	1178.98	705.77
	(1923.94)	(8400.14)	(2135.62)	(12555.75)
Export price index	-59.34**	-26.31**	-20.19	-13.49
	(28.61)	(9.11)	(20.56)	(22.56)
N	41	41	41	41
R-squared	0.90	0.60	0.78	0.79
Time dummies	Yes	Yes	Yes	Yes
Indiv. effect	Random	Fixed	Random	Fixed

* p<0.1, ** p<0.05, *** p<0.01. Cluster-robust standard errors reported.

Table 3.7: Panel OLS augmented with income and export price index

two columns of each table use lagged values of the independent variables as controls, while the last two incorporate exogenous instruments. I report tests for under- and over-identification in both tables. Also reported are tests for the exclusion restrictions for a subset of the instruments used, where the null hypothesis is that both the smaller set of included instruments and the subset of instruments being tested for exogeneity are valid. The IV diagnostic tests do not indicate an obvious misspecification of the instrument set. Details on the instruments used are given in Appendix F.

Tables 3.8 and 3.9 show that when instrumenting for levels of log GDP per capita and the export value/price indices in the regression models, the conditional correlation between democracy and the disparity in trade policy expenditure remains robust. The negative correlation between each country's Polity score and its tariff-subsidy expenditure difference, and the positive correlation between the difference and the Freedom in the World index, shows that dictatorships levy higher values of import duties and tariffs compared to export subsidy payments than do democracies. This conditional correlation is in the direction predicted by the model in Section 3.3, and suggests a systematic pattern between domestic political institutions and a country's choice between alternative interventionist trade policies.

	REG13	REG14	REG15	REG16
Log GDP per capita	8140.60***	8548.57***	24903.74***	28630.83***
	(1733.75)	(1716.11)	(9206.85)	(10583.84)
Export value index	-16.75**	-18.33**	10.78	36.01
	(8.34)	(7.95)	(33.43)	(52.79)
polity2	-40.87*		-177.80*	
	(21.30)		(104.16)	
findex		491.65***		1375.91**
		(128.93)		(563.81)
N	98	101	128	131
Root MSE	628.09	615.25	1268.33	1458.09
p-value	0.00	0.00	0.05	0.05
Under-i.d. p	0.00	0.00	0.08	0.09
Over-i.d. p	0.72	0.84	0.20	0.18
Exog. test p	0.72	0.84	0.37	0.18
Time dummies	Yes	Yes	Yes	Yes
Indiv. effect	Fixed	Fixed	Fixed	Fixed

* p<0.1, ** p<0.05, *** p<0.01. Cluster-robust standard errors reported.

Table 3.8: Panel IV with income and export value index

	REG17	REG18	REG19	REG20
Log GDP per capita	41844.11**	95601.94*	11931.26**	12705.78
	(18255.76)	(54011.25)	(5587.18)	(11313.45)
Export price index	1.76	17.48	-28.17***	-34.15**
	(21.76)	(38.58)	(5.80)	(14.07)
polity2	-1515.13***		-1388.66***	
	(229.40)		(189.38)	
findex		4302.88***		3521.30***
		(1282.76)		(795.09)
N	41	41	41	41
Root MSE	529.16	1033.00	413.10	619.27
p-value	0.00	0.08	0.00	0.01
Under-i.d. p	0.07	0.06	0.05	0.07
Over-i.d. p	0.94	0.96	0.23	0.42
Exog. test p	0.90	0.93	0.46	0.78
Time dummies	Yes	Yes	Yes	Yes
Indiv. effect	Fixed	Fixed	Fixed	Fixed

* p<0.1, ** p<0.05, *** p<0.01. Cluster-robust standard errors reported.

Table 3.9: Panel IV with income and export price index

3.4.3 Robustness Checks

As a further robustness check, I perform a similar exercise to the first chapter in this dissertation, using military sophistication as a time-varying proxy for the exogenous value of office-holding. Political candidates may value holding political office more if the government has greater international influence. I proxy for international influence with military sophistication, defined as the log ratio of military expenditure per personnel. Military expenditure data is taken from the World Development Indicators and is expressed in year 2000 U.S. dollars. Total armed forces personnel is also from the WDI. I augment the IV specification (instrumenting for log GDP per capita and the export price/value indices) with this proxy for the value of holding office, treating military sophistication as exogenous to the difference between tariff and subsidy expenditure. Table 3.10 shows that incorporating this additional control does not change the qualitative prediction regarding democracy and the disparity between trade policy expenditure.

	REG21	REG22	REG23	REG24
Log GDP per capita	8004.59***	8997.93***	31865.02**	53463.48
	(2168.16)	(2121.98)	(16136.31)	(36572.19)
Military sophistication	115.72	-37.57	-253.56	3417.16
	(482.75)	(445.74)	(1520.20)	(3191.82)
Export value index	-18.53**	-19.72***		
	(8.30)	(7.62)		
Export price index			9.47	-27.68
			(19.71)	(35.45)
findex		489.13***		3470.87***
		(132.32)		(1164.58)
polity2	-43.77**		-1497.86***	
	(21.30)		(229.13)	
N	97	97	41	41
Root MSE	629.14	592.76	488.84	754.51
p-value	0.00	0.00	0.00	0.13
Under-i.d. p	0.00	0.00	0.03	0.07
Over-i.d. p	0.75	0.82	0.84	0.74
Exog. test p	0.63	0.72	0.78	0.64
Time dummies	Yes	Yes	Yes	Yes
Indiv. effect	Fixed	Fixed	Fixed	Fixed

* p<0.1, ** p<0.05, *** p<0.01. Cluster-robust standard errors reported.

Table 3.10: Panel IV with proxy for value of holding office

Additionally, it should be noted that while Y_H in the model reflects citizens' income endowments, it also reflects the volume of exports out of the economy. The results in the above tables can be shown to be robust to controlling for export volumes without the need to gather additional data. Changes in the export value index can reflect fluctuations in a country's export prices as well as its export volume. Incorporating both the export price index and the export value index in the regression will allow us to examine the effect of a change in the democracy index on the difference in tariff and subsidy expenditure, keeping export prices and volumes constant. This is what I do in Table 3.11. For given levels of export prices and volumes, an increase in the Freedom in the World index increases tariff expenditure relative to export subsidy expenditure. A decrease in the Polity index has the same effect. This illustrates that compared to democratic governments, authoritarian regimes collect higher levels of import tariffs relative to export subsidy expenditure.

	REG25	REG26	REG27	REG28
Export price index	43.94*	9.88	43.04	10.36
	(25.42)	(13.97)	(28.28)	(19.47)
Export value index	-74.90***	-41.64***	-74.67***	-40.53**
	(21.89)	(13.52)	(22.95)	(16.28)
Military sophistication			9.90	-62.23
			(1256.83)	(769.96)
findex	1664.55*		1673.14*	
	(878.05)		(877.35)	
polity2		-975.38***		-983.21***
		(202.80)		(216.18)
N	41	41	41	41
Root MSE	450.17	356.03	450.41	356.26
p-value	0.00	0.00	0.00	0.00
Under-i.d. p	0.07	0.09	0.09	0.09
Over-i.d. p	0.12	0.18	0.12	0.16
Exog. test p	0.79	0.84	0.80	0.80
Time dummies	Yes	Yes	Yes	Yes
Indiv. effect	Fixed	Fixed	Fixed	Fixed

* p<0.1, ** p<0.05, *** p<0.01. Cluster-robust standard errors reported.

Table 3.11: Panel IV controlling for export volume instead of income

Finally, Table 3.12 builds on Table 3.11, with the addition of the import value index as a control. Data on the import value index were from the World Development

Indicators. Fluctuations in this index reflect fluctuations in both the price and volume of imports. While the previous regression specifications controlled for the determinants of subsidy expenditure sY_H suggested by the model, tariff expenditure tP_FC_F varies with the prices and quantity of imports, where t is given by equation (3.20). Controlling for the import value index in the regression allows variations in tariff revenue to reflect changes in policy rather than changes in the prices and quantity of imports. The first and second column of Table 3.12 treat the import value index as included (exogenous) instruments, while the third and fourth column instrument for the import value index. Data on the instruments used are in Appendix F. The table shows that the previous results are robust to the inclusion of this additional regressor. For given values of the import value index, export price index, export quantities, and military sophistication, an increase in the Polity index (i.e., an increase in the level of democracy) is significantly associated with lower tariff revenues relative to export subsidy expenditures.

	REG29	REG30	REG31	REG32
Export price index	-1.05	14.48	1.54	21.97
	(27.97)	(27.72)	(17.91)	(24.33)
Export value index	-50.69**	-46.85**	-29.11***	-22.08
	(21.52)	(23.87)	(9.79)	(15.18)
Import value index	18.89	2.23	-0.10	-35.01
	(40.14)	(38.43)	(27.14)	(35.00)
Military sophistication		-397.08		623.10
		(943.66)		(766.60)
polity2	-1049.85***	-902.29***	-1052.27***	-736.12**
	(343.88)	(326.34)	(258.59)	(314.21)
N	41	41	41	41
Root MSE	361.86	360.51	364.20	370.18
p-value	0.00	0.00	0.00	0.00
Under-i.d. p	0.07	0.08	0.07	0.07
Over-i.d. p	0.18	0.14	0.42	0.27
Exog. test p	0.94	0.74	0.29	0.54
Time dummies	Yes	Yes	Yes	Yes
Indiv. effect	Fixed	Fixed	Fixed	Fixed

* p<0.1, ** p<0.05, *** p<0.01. Cluster-robust standard errors reported.

Table 3.12: Panel IV controlling for import value index

3.5 Conclusions

When policymakers are motivated by rent-seeking behavior, trade policies may be more biased towards import tariffs than export subsidies. Democratic institutions prevailing in a country can reduce the extent of this bias. This paper has presented a theoretical model and empirical evidence in support of this finding. While most papers in the trade literature recast import tariffs as a tool to protect domestic import-competing industries, thinking about tariffs as a means of enriching opportunistic politicians adds to our understanding of why (supposedly protectionist) trade policies are systematically biased against trade—i.e., in favor of import tariffs rather than export subsidies—across countries (Rodrik 1995). The tight interlinkage between political institutions and preferred interventionist trade policies, combined with the gradual pace with which political institutions change, indicate the magnitude of the challenges in negotiating free trade agreements amongst countries with different attitudes toward democracy and explain why some trade arrangements find it necessary to address non-trade-related, political concerns (Aaronson 2007). In lieu of achieving perfect political convergence, international trade organizations consisting of members with widely disparate political institutions may find it easier to address trade barriers which cannot be used for political rent-seeking, prompting arrangements like the WTO to be stricter on export subsidies than import tariffs (Potipiti 2012). For further research, it may be promising to investigate and explain whether this systematic political asymmetry in trade intervention exists for other policies, such as export taxes versus tariffs, as well as analyze the welfare effects of the interaction between different political regimes in a multilateral free trade agreement.

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Appendix F

Details on instruments used

Model	Instrument Set	Subset of instrments tested for exogeneity
REG13	1st through 2nd lag of log GDP per capita 1st through 2nd lag of export value index	1st through 2nd lag of log GDP per capita
REG14	1st through 2nd lag of log GDP per capita 1st through 2nd lag of export price index	1st through 2nd lag of log GDP per capita
REG15	Current through 3rd lag of fertility rate Current and lagged openness	Current and lagged openness
REG16	Current and lagged fertility rate Lag of openness	Lag of openness
REG17	3rd through 5th lag of log GDP per capita 3rd through 5th lag of export price index	3rd through 5th lag of log GDP per capita
REG18	3rd through 5th lag of log GDP per capita 2nd through 5th lag of export price index	3rd through 5th lag of log GDP per capita
REG19	Current through 3rd lag of fertility 2nd through 5th lag of openness	2nd through 5th lag of openness
REG20	Current through 3rd lag of fertility 2nd through 5th lag of openness	2nd through 5th lag of openness
REG21	1st through 3rd lag of log GDP per capita 1st through 2nd lag of export value index	1st to 2nd lag of log GDP per capita
REG22	1st through 3rd lag of log GDP per capita 1st through 2nd lag of export value index	1st to 2nd lag of log GDP per capita

Model	Instrument Set	Subset of instrments tested for exogeneity
REG23	2nd through 5th lag of log GDP per capita 3rd to 5th lag of export price index	2nd to 5th lag of log GDP per capita
REG24	3rd through 5th lag of log GDP per capita 2nd to 5th lag of export price index	3rd to 5th lag of log GDP per capita
REG25	Current and lag import share of GDP Current and lag fertility rate Lag of openness	Current and lag import share of GDP
REG26	Current and lag import share of GDP Current and lag fertility rate Lag of openness	Current and lag import share of GDP
REG27	Current and lag import share of GDP Current and lag fertility rate Lag of openness	Current and lag import share of GDP
REG28	Current and lag import share of GDP Current and lag fertility rate Lag of openness	Current and lag import share of GDP
REG29	Current and lag import share of GDP Current and lag fertility Lag of openness	Import share of GDP
REG30	Current and lag import share of GDP Current and lag fertility Lag of openness	Import share of GDP
REG31	Current through third lag of: import share of GDP, export share of GDP, and fertility rate	Import share of GDP Export share of GDP
REG32	Current through 2nd lag of: import share of GDP, export share of GDP, and fertility rate	Import share of GDP Export share of GDP

Openness measured as trade (sum of exports and imports) share of GDP.