Price Fluctuations and Political Conflict

John O'Trakoun *

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Abstract

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. 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.

Keywords: Political conflict; Commodity prices; Uncertainty; Political economy.
JEL Classification Numbers: P480, P420, O500.

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**

*Boston College, Department of Economics, 140 Commonwealth Avenue, Chestnut Hill MA 02467. otrakoun@bc.edu

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 3, I lay out a simple, two-period small open economy model of political conflict. Section 4 describes the data and presents empirical results supporting the model, and Section 5 concludes.

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 3.

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.

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.

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:

$$E_0[U_1] = E_0 \left[\ln C_1 + \beta \ \ln C_2\right] \tag{1}$$

subject to the budget constraint

$$C_t \leq P_t(1-\tau)m \tag{2}$$

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:

$$\pi_t = \begin{cases} 0, & \text{if } U_t < \overline{\omega_t} \\ 1, & \text{if } U_t \geq \overline{\omega_t} \end{cases} \tag{3}$$

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:

$$E_t\left[v_D\right] = \ln r + \pi_t \ln R \tag{4}$$

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

$$r = \tau m \tag{5}$$

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.

$$\ln r + \ln R \geq \ln m \quad \Rightarrow \quad r \geq m/R \tag{6}$$

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.

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 (3) holds as an equality. Combining (3) with (1) and the distribution of world prices yields the condition:

$$\overline{\omega_1} = (1+\beta)\left[\ln(1-\tau)m + \mu\right] \tag{7}$$

Applying the dictator budget constraint (5), it is possible to derive one of the conditions for the equilibrium level of rents:

$$r = m - \exp\left(\frac{\overline{\omega_1}}{1+\beta} - \mu\right) \tag{8}$$

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 (6) hold with equality.

$$\ln r + \ln R = \ln m \;\Rightarrow\; r = \frac{m}{R} \tag{9}$$

Equations (8) and (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:

$$\begin{aligned} U_1 &= \ln P_1(1-\tau)m + \beta E_1\left[\ln P_2(1-\tau)m\right] \\ &= \ln P_1(1-\tau)m + \beta\left[\ln(1-\tau)m + \mu\right] \\ &= \ln P_1(1-\tau)m + \beta\left[\ln(1-\tau)m + \ln E_1[P_2] - \tfrac{\sigma^2}{2}\right] \end{aligned} \tag{10}$$

The last two lines in (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 (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 (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.

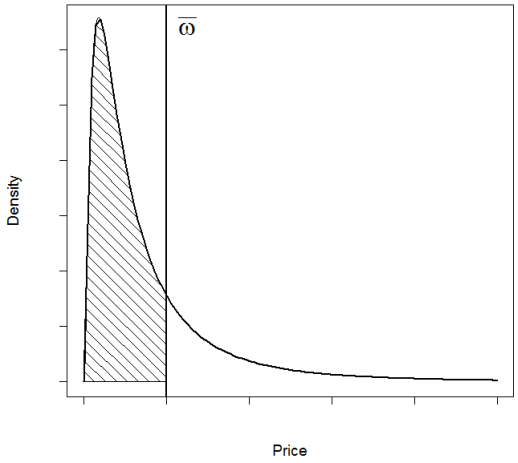
3.4 Comparative Statics

Using (8) and (9), the solution for the reservation level in period 1 is:

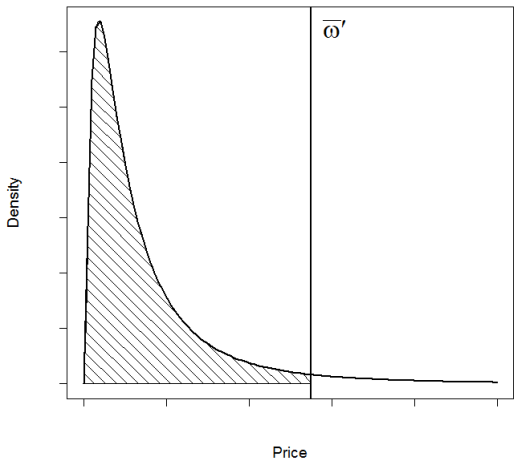
$$\overline{\omega_1} = (1+\beta)\left\{\ln m + \ln\left(\frac{R-1}{R}\right) - \mu\right\} \tag{11}$$

Taking the derivative of (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 (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 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 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.

Similarly, taking the derivative of (11) with respect to the exogenous value of holding power, R, yields: $\dfrac{\partial\overline{\omega_1}}{\partial R} = \dfrac{1+\beta}{R(R-1)}$. 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



(a) Lower m



(b) Higher m

Figure 1: Effect of increase in m

citizens; combining equilibrium conditions (5) and (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.

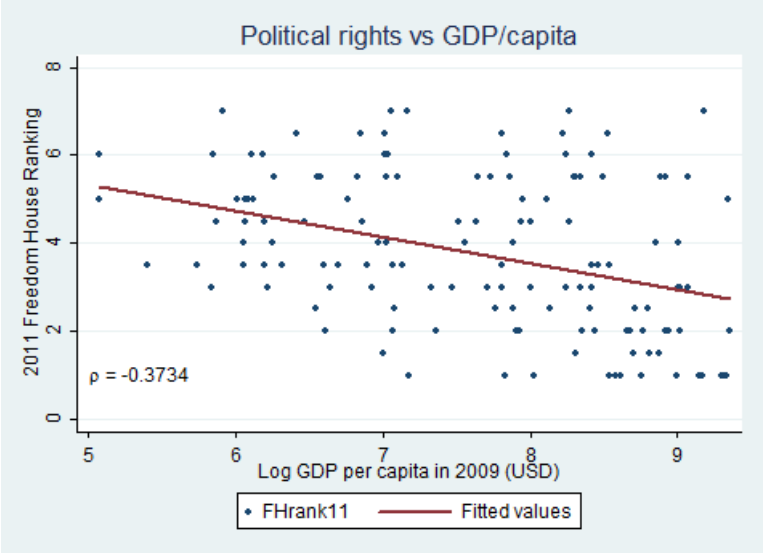
These results find some support in the data. In the graphs below, I use the 2011 Freedom in the World survey's combined average rating of political rights and civil liberties[1] to proxy for whether or not a country has overthrown a dictator. Rating is done on a scale from 1 to 7, with lower scores indicating 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 (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. In Figure (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

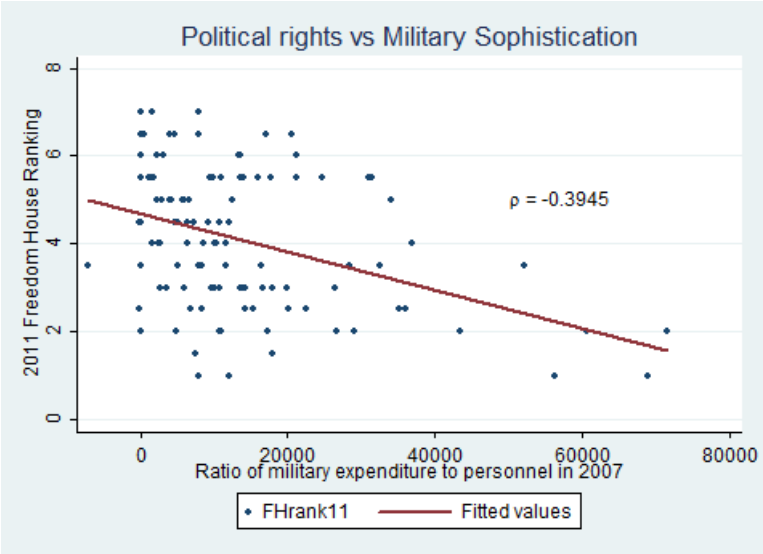
[1] Published by Freedom House (see appendix).

[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.

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

4 Empirical Results

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:

$$\text{conflict}_t = \begin{cases} 1, & \text{if conflict in period t} \\ 0, & \text{if no conflict in period t} \end{cases} \tag{12}$$

Descriptive statistics for the conflict variable are given in Table 1.

To measure prices, I use international commodity price data from the *World Bank*

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

Table 1: Summary statistics for conflict variable, quarterly frequency

GEM Commodities database. (Bank 2011) The data used are given at monthly 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 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

[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 2: Summary statistics for prices and uncertainty series, quarterly frequency

are dropped from the sample, as these countries are most likely to violate the small 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 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 3.

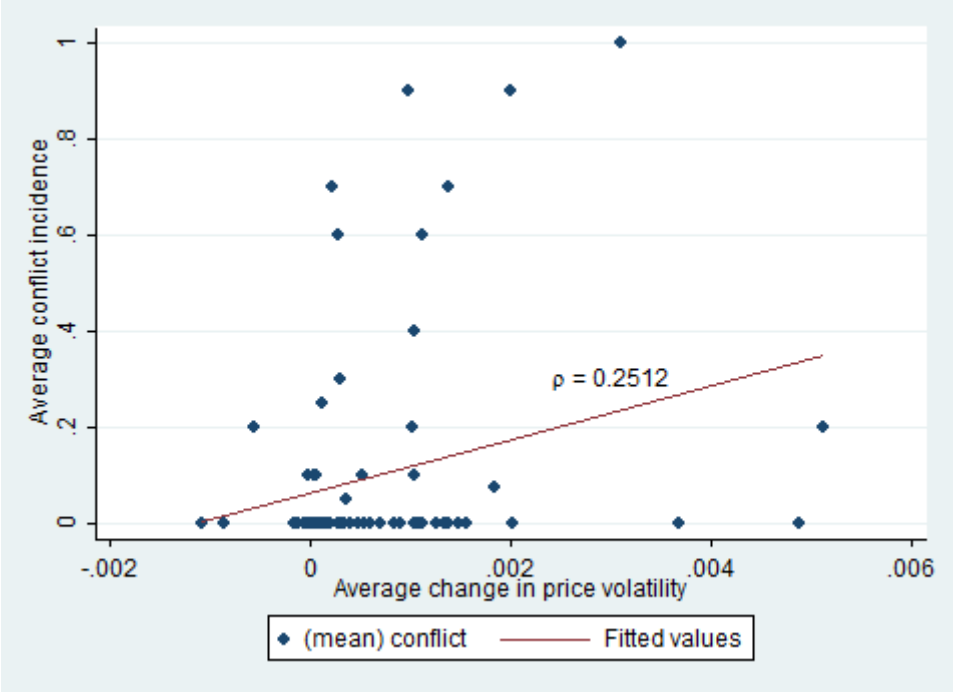


Figure 3: Average conflict incidence versus average change in uncertainty, 1960s

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 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,

$$\ln P_t = \ln P_{t-1} + \sigma_{t-1}\eta_t \tag{13}$$

$$\sigma_t = \sigma_{t-1} + \epsilon_t \tag{14}$$

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:

$$E_0[U_1] = \overline{\omega_1} = \ln P_0 + \beta \ln E_0[P_1] - \frac{\beta\sigma_0^2}{2} + (1+\beta)\ln\{(1+\tau)m\} \tag{15}$$

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

$$U_1 = \ln P_1 + \beta \ln E_1[P_2] - \frac{\beta\sigma_1^2}{2} + (1+\beta)\ln\{(1+\tau)m\} \tag{16}$$

Conflict occurs when the difference between (16) and (15) is less than zero:

$$\text{conflict}_1 = \begin{cases} 1, & \text{if } \Delta \ln P_1 + \beta\left(\ln E_1[P_2] - \ln E_0[P_1]\right) - \frac{\beta}{2}\Delta\sigma_1^2 < 0 \\ 0, & \text{otherwise} \end{cases} \tag{17}$$

Equation (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

[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.

conflict. Equation (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, including changes in price level and uncertainty as regressors.

4.2.1 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 3: Effect of price level and volatility on conflict incidence

Table 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

[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. $P = 0.4813$.

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 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 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 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 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 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 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 4 are largely similar to those in Table 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 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 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 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 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 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 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

[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.

is to impute higher frequency data from more the complete data series available at the annual frequency. This is what I do in Table 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 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 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 6: Effect of price level, volatility, and income on conflict incidence

4.3 Robustness Checks

4.3.1 Alternative Measures of Volatility

In Tables 7 & 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 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 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.

Table 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 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

	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 7: Estimation using price variance to measure price uncertainty

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.

4.3.2 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

	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)
ΔVol_2PI_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)
$\Delta rain_t$		0.015				
		(0.015)				
$\Delta reserves_t$			-0.033	-0.057	-0.033	
			(0.040)	(0.198)	(0.110)	
$\Delta income_t$						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 8: Estimation using price range to measure price uncertainty

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 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 proportion-

ally interpolated measure of income per capita in the fourth through sixth columns of Table 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 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 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 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.

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:

$$onset_{1t} = \begin{cases} 0, & \text{if conflict}_t = 0 \\ 1, & \text{if conflict}_t = 1 \& \text{ conflict}_{t-1} = 0 \\ 1, & \text{if conflict}_t = 1 \& \text{ conflict}_{t-1} = 1 \end{cases} \tag{18}$$

then conflict outbreak and conflict incidence would exactly equal one another; conflict

	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 10: Effect of price uncertainty (variance and s.d.) on onset of 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:

$$onset_{2t} = \begin{cases} 0, & \text{if conflict}_t = 0 \\ 1, & \text{if conflict}_t = 1 \& \text{ conflict}_{t-1} = 0 \cdot \\ 0, & \text{if conflict}_t = 1 \& \text{ conflict}_{t-1} = 1 \end{cases} \tag{19}$$

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 11 replicates the estimation of the models in Table 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 11: Effect of price uncertainty (variance and s.d.) on modified onset measure

4.3.3 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 12 and 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 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 12. Results are qualitatively similar when examining conflict outbreak, as seen in Table 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 12 and 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 13, which uses the fourth through the eighth lags of changes in rainfall to instrument for changes in GDP per capita.

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

	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)				
Δreserves$_t$			-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 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)
ΔVol$_1$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 13: Price volatility and conflict onset, using year 1996 net export shares as weights

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.

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

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 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:

$$\begin{aligned} U_1 &= \log \tfrac{(1-\tau)m}{\tilde{P}_1} + \beta E_1 \left[\log \tfrac{(1-\tau)m}{\tilde{P}_2} \right] \\ &= \log \tfrac{(1-\tau)m}{\tilde{P}_1} + \beta \left[\log (1-\tau)m - \mu \right] \\ &= \log \tfrac{(1-\tau)m}{\tilde{P}_1} + \beta \left[\log (1-\tau)m + \log E_1[\tfrac{1}{\tilde{P}_2}] - \tfrac{\sigma^2}{2} \right] \end{aligned} \tag{20}$$

where the last two lines in (20) 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 (20), 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 (20), 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 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 4. Summary statistics given in Table 15 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 16 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 15: Summary statistics for net import share weighted price and uncertainty series, quarterly frequency

	IV-LPM-FE	LPM-FE	LOGIT-RE	LOGIT-FE	PROBIT-RE
$\Delta \text{PI}_{imp,t}$	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 16: Effect of net import-weighted price level, volatility, and income on conflict incidence

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 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 17 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 17: Price volatility and conflict likelihood, omitting China and India

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