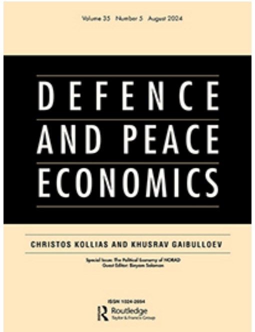

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Asymmetric Effects of Positive and Negative Commodity Price Shocks During Civil Wars

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ABSTRACT
Economic shocks, including shocks to commodity export prices, have often been related to civil conflict. However, the literature usually assumes that commodity prices have linear effects. In this paper, we hypothesize that negative and positive commodity export price shocks can have asymmetric effects: while negative shocks are likely to increase conflict, positive shocks can have a smaller conflict-reducing effect or even increase conflict. We test this hypothesis in a dataset for battle deaths during ongoing civil wars. Consistent with the hypothesis, negative price shocks increase battle deaths, but positive price shocks also have a positive effect. The positive effect is concentrated and becomes significant in countries with weak pre-conflict institutions that experience positive price shocks to fossil fuels. We conclude that a combination of institutional reforms, economic diversification, and price stabilization might reduce battle deaths in fuel-dependent countries with weak institutions.

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Introduction

Civil wars have many adverse effects, including casualties, population displacement, health and education losses, physical capital losses, and environmental destruction. Although the causes of civil wars remain imperfectly understood, economic shocks, including shocks to commodity export prices, might play an important role (Blattman and Miguel 2010; Blair, Christensen, and Rudkin 2021). However, the literature usually assumes that commodity prices have a linear effect. The possibility that commodity prices can have asymmetric effects has rarely been explored and represents a possible gap in the literature on commodity prices and civil conflict. In this paper, we hypothesize that negative and positive commodity export price shocks can have asymmetric effects. First, we develop a theoretical framework that explains or summarizes (Moshiri 2015) five possible causes of asymmetry and explains why we might expect weak institutions and extractive-industry dependence to increase the degree of asymmetry. Second, we test the hypothesis in a dataset of battle deaths during ongoing civil wars from 1970-2008. Consistent with this hypothesis, negative price shocks increase battle deaths and positive shocks also have a positive effect. The positive effect is concentrated and becomes statistically significant in countries that have weak pre-conflict institutions (defined as unconstrained or undemocratic governments) and experience a positive price shock to fossil fuels. We conclude that a combination of institutional reforms, economic

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diversification, and price stabilization (Van der Ploeg and Poelhekke 2009; Frankel 2017) might reduce battle deaths in fuel-dependent countries with weak institutions.

Although we fully develop the theoretical framework in Section 3, why, intuitively, might we expect positive and negative commodity price shocks to have asymmetric conflict effects? Why are the asymmetric effects largely explained by countries with weak institutions that experience positive fuel price shocks? One of the possible explanations builds on Moshiri (2015), who argues that governments tend to play an outsized economic role in oil exporters with weak institutions. In such countries, governments collect large royalties from oil and conduct procyclical fiscal policy. Positive oil price shocks increase government spending on unproductive social programs and investment projects. When the price falls, the government continues to support these programs and projects to prevent social unrest, forcing budget cuts elsewhere. Moshiri (2015) shows that in developing countries with weak institutions, negative oil price shocks cause large decreases in economic growth. In contrast, positive shocks do not increase growth. Given the robust finding in the conflict literature that economic growth tends to reduce conflict (Blattman and Miguel 2010), we should expect a negative price shock to increase conflict while a positive shock should not necessarily decrease conflict. Moreover, weak institutions should increase the appropriability of the price gain (Boschini, Pettersson, and Roine 2007), which alone should increase conflict after a positive price shock. Weak institutions should also weaken public and private coping and insurance mechanisms after negative price shocks, such as the ability to accumulate and sell household assets, credit access, unemployment insurance, and public employment programs. This should exacerbate household distress and cause more conflict after a negative price shock. Finally, dictators, military juntas, and other unconstrained governments are less likely to be concerned with civilian casualties and human rights abuses (Davenport and Armstrong 2004; Bellamy 2011). After a negative price shock, they can respond to increased insurgency activity with indiscriminate government violence, such as shelling or burning villages. After a positive price shock, they can more freely increase military spending and counterinsurgency attacks regardless of the death toll.

In the remainder of the paper, Section 2 relates the paper to the literature. Section 3 develops our theoretical framework that explains why positive and negative commodity price shocks can have asymmetric effects on conflict. Section 4 presents the data and empirical specification. Section 5 presents the results. Section 6 contains robustness checks. Section 7 concludes the paper.

Existing Literature

The paper is related to the literature on natural resources and conflict (Ross 2004; Lujala 2009, 2010; Koubi et al. 2014; Nillesen and Bulte 2014), the effects of commodity prices on households and the macroeconomy (Lederman and Porto 2016; Gruss and Kebhaj 2019), and, particularly, the literature relating commodity prices to conflict (Dube and Vargas 2013; Blair, Christensen, and Rudkin 2021).[1] According to this literature, higher commodity prices can increase the opportunity cost of fighting and decrease conflict (Miguel, Satyanath, and Sergenti 2004; Blattman and Miguel 2010; Dube and Vargas 2013). However, they can also increase the incentive to appropriate natural resources (Boschini, Pettersson, and Roine 2007, 2013; Dube and Vargas 2013) and help to finance existing conflict (Keen 2000; Le Billon 2001; Bannon and Collier 2003; Janus 2012). The appropriation and financing effects may be stronger in economies with weak institutions because of contestable property rights and the lack of alternative financing mechanisms (Bannon and Collier 2003; Mehlum, Moene, and Torvik 2006; Global Witness 2009), when the resource is geographically concentrated (Maystadt et al. 2014; Berman et al. 2017; Romsom 2022), and when resource extraction is capital intensive (Dal Bó and Dal Bó 2011; Dube and Vargas 2013). For example, Brückner and Ciccone (2010), Janus and Riera-Crichton (2015), and Berman et al. (2017) estimate the effect of commodity prices on the risk of onset or incidence of civil conflict. Bazzi and Blattman (2014) estimate both onsets and the intensity of ongoing civil conflicts. In this paper, while we follow Bazzi and Blattman's (2014) focus on estimating the intensity of ongoing conflicts, we develop a theoretical framework and test for asymmetric effects of commodity prices. Additionally, we estimate the role of pre-conflict institutional quality as measured by government accountability.[2]

The paper is also related to a small literature that tests for asymmetric effects of commodity price shocks on economic growth in commodity exporters. Moshiri (2015) estimates the effects of oil price shocks on nine oil exporters and find that, in developing countries with weak institutions, negative oil price shocks cause large declines in government revenues and economic growth. In contrast, positive shocks do not increase economic growth. In a study of Iran, Farzanegan and Markwardt (2009) find that negative oil price shocks increase inflation and decrease economic growth and short-run government spending. Positive price shocks increase economic growth, also increase inflation, and increase medium-run (but not short-run) government spending. Emami and Adibpour (2012) find that negative oil price shocks in Iran deeply decrease output growth while positive shocks only have a moderate positive effect. The authors attribute this result to the resource curse. Although we estimate the intensity of conflict rather than economic growth, the disappointing effects of positive oil price shocks in these papers compared to the harm done by negative shocks are consistent with our results that positive price shocks fail to decrease battle deaths (Blattman and Miguel 2010).

Theoretical Framework

In this section, we explain why we might expect positive and negative economic shocks to have asymmetric effects on conflict. First, Moshiri (2015) studies a sample of developed and developing country oil exporters with different degrees of institutional quality. The author argues that in developing-economy oil exporters with weak institutions, governments tend to play an outsized economic role, collect large royalties from oil, and conduct procyclical fiscal policy. Positive oil price shocks increase government spending on unproductive social programs and investment projects. The burden from these fiscal obligations accumulates over time. When prices fall, the government must continue to support these programs and projects to prevent social unrest, which forces budget cuts elsewhere. Moshiri (2015) estimates that in developing oil exporters with weak institutions, negative oil price shocks cause large decreases in government revenues and economic growth. In contrast, positive shocks do not increase growth. In the context of this paper, therefore, and given that economic growth tends to reduce conflict (Blattman and Miguel 2010), we might expect that a positive shock to oil prices should insignificantly decrease conflict. Since the conflict 'prize' should still increase because of the positive price shock, we could even observe an increase in conflict. At time same time, when prices fall, the decline in economic growth decline and the budget cuts might increase the citizens' opportunity cost of fighting, frustration, desperation, and violence. If weak institutions also weaken public and private coping and insurance mechanisms, such as the ability to accumulate and market household assets, unemployment insurance, public programs, and private credit access, this should further increase household distress and conflict.

Second, wages may be downwardly rigid in developing countries (Schmitt-Grohé and Uribe 2016; Kaur 2019). In that case, a decline in export prices might cause the unemployment rate to spike or erode the government's safety net, sending workers into the conflict sector. Again, if the government is relatively weak, corrupt, or ineffective (Gould and Winters 2012; Amundsen 2014), or the private economy is weakly developed and fails to provide coping mechanisms like credit and insurance, the consumption decline may be larger and conflict should increase more. In contrast, when the export price increases, the fact that wages are upwardly flexible implies that wages will increase. The increase in wages tends to limit the increase in hiring, thus, pulling relatively few workers out of the conflict sector in the short run. This mechanism is reinforced if weak institutions increase appropriability (Boschini, Pettersson, and Roine 2007).

Third, even if wages are flexible or most individuals are self-employed, the labor supply curve may be flat at low employment levels and steep at high employment levels as households approach their time endowment and the marginal utility of leisure relative to consumption becomes high (Dessing 2002; Cuaresma 2003; Ocampo 2005). This has the same effect as downward nominal wage rigidity: outward shifts in labor demand mainly affect wages whereas inward shifts affect employment. Fourth, loss aversion (Thaler and Johnson 1990; Jervis 1992) or the need to prevent income from falling below a threshold that would trigger a poverty trap might encourage risk-taking, including potentially the uptake of arms, after negative price shocks. Weak institutions that undermine coping and insurance mechanisms should encourage individuals to take more desperate measures after a negative price shock and again increase appropriability after a positive price shock.

Fifth, if we conceive of the civil war as a dynamic game between the government and armed groups, a temporary economic and government revenue decline might create a window of opportunity for rebel attacks. Conversely, temporary economic and revenue increases might create a window for government attacks (Chassang and I Miquel 2009). Fearon (2004) presents a model where regional autonomy agreements may be unreachable when fluctuations in the government's strength – which might be induced by positive and negative commodity price and tax-revenue shocks – undermine its ability to commit. This commitment problem is more likely to lead to conflict when the government has an enduring political or economic interest in expanding into the region and when either the government or the rebels earn income despite the conflict (e.g. through natural resource exports). Self-interested governments or governments that maximize the interest of only specific population segments (such as the military or specific ethnic groups) should be more likely to be observed in countries with weak institutions. Thus, the commitment problem should be greater, causing more conflict.

The asymmetric effects of positive and negative price shocks might be larger in countries that are characterized not only by weak institutions but dependence on extractible as opposed to agricultural natural resources: First, the literature on the natural resource curse suggests that the combination of weak institutions and extractible natural resources might be particularly detrimental to economic development and institutional quality (Boschini, Pettersson, and Roine 2007, 2013; Van der Ploeg 2011; Dube and Vargas 2013; Waldner and Smith 2015; Cassidy 2019). If countries that depend on extractible resources, such as oil and minerals, are characterized by poorer households, weak property rights, less market development, and less state capacity and accountability, they should have weaker coping mechanisms after negative price shocks and households may be more likely to turn to conflict. Second, extractive compared to agricultural income may be easier to tax or the government might own oil and mineral companies directly. In that case, price declines will have a larger impact on state capacity, emboldening insurgents. Due to a lack of accountability and alternative revenue sources, positive extractible price shocks may be spent on the military and violence-intensive counter-insurgency (Manz 1988; Paul, Clarke, and Grill 2010; Bellamy 2011) rather than on reducing violence, public goods, and welfare programs. Thus, the combination of extractible resources and weak institutions might exacerbate the adverse conflict effects of positive and negative price shocks.

Altogether, we test the following hypotheses:

H0a: *Positive and negative shocks to commodity export prices have asymmetric effects*

H1a: *Positive and negative shocks to commodity export prices have symmetric effects*

H0b: *Positive and negative shocks to commodity export prices have asymmetric effects when institutions are weak*

H1b: *Positive and negative shocks to commodity export prices have symmetric effects when institutions are weak*

H0c: *Positive and negative shocks to extractive commodity export prices have asymmetric effects when institutions are weak*

H1c: *Positive and negative shocks to extractive commodity export prices have symmetric effects when institutions are weak*

Data and Specification

We estimate an annual panel of countries with ongoing civil wars from 1970 to 2008. The start date reflects that the trade data we need to construct the CTOT index is scarce before 1970. The battle deaths data ends in 2008 (Lacina and Gleditsch 2005).[3]

Battle Deaths

We use data for battle-related fatalities in internal and internationalized internal armed conflicts from Lacina and Gleditsch (2005). Armed conflict, internal and internationalized armed conflict, and battle-related fatalities are defined as (Gleditsch et al. 2002; Themnér and Wallensteen 2011, Codebook for the UCDP/PRIO Armed Conflict Dataset, Version 4, 1, 9; Lacina and Gleditsch 2005, Battle Deaths Dataset, Codebook for Version 3, 2009, 2):

> [An armed conflict is] a contested incompatibility that concerns government or territory or both where the use of armed force between two parties results in at least 25 battle-related deaths. Of these two parties, at least one is the government of a state.
>
> Internal armed conflict occurs between the government of a state and one or more internal opposition group(s) without intervention from other states.
>
> Internationalized internal armed conflict occurs between the government of a state and one or more internal opposition group(s) with intervention from other states (secondary parties) on one or both sides.
>
> [Battle-related fatalities are] those deaths caused by the warring parties that can be directly related to combat over the contested incompatibility. This includes traditional battlefield fighting, guerrilla activities (e.g. hit-and-run attacks/ambushes) and all kinds of bombardments of military bases, cities and villages etc. Urban warfare (bombs, explosions, and assassinations) does not resemble what happens on a battlefield, but such deaths are considered to be battle-related. The target for the attacks is either the military forces or representatives for the parties, though there is often substantial collateral damage in the form of civilians being killed in the crossfire, indiscriminate bombings, etc. All fatalities – military as well as civilian – incurred in such situations are counted as battle-related deaths.

Since the number of battle-related fatalities is surrounded by significant uncertainty, Lacina and Gleditsch (2005) provide 'low' and 'high' estimates for each conflict year. Additionally, they provide a 'best' estimate for approximately 70% of the observations. Since the data is at the country-year-conflict level, we add the low, high, and best estimates within country-years to get a country-year panel of low, high, and best estimates.

Our main empirical measure is the best country-year estimate unless it does not exist. If it does not exist, we use an 'imputed' best estimate. The imputed best estimate is the sum across ongoing conflicts within a country-year of the best conflict-specific estimate or if it does not exist, the average of the low and high estimates. This methodology follows Bazzi and Blattman's (2014). If we alternatively dropped the country-years with missing best estimates, we might be dropping a selected sample. For example, countries with multiple ongoing conflicts – making missing data more likely – may have weak institutions. In any case, as we show below, the qualitative results are robust to using the low or high instead of the 'best' and 'imputed best' estimates. They are also robust to using only battle-death estimates derived from year-specific sources in Lacina and Gleditsch (2005), which excludes most of the imputed observations.

Commodity Terms of Trade

The CTOT index is an index of GDP-weighted commodity export relative to import prices proposed by Ricci, Milesi-Ferretti, and Lee (2008) and Spatafora and Tytell (2009)[4] and provided by Janus and Riera-Crichton (2015). The appendix to Janus and Riera-Crichton (2015) further details how the data is constructed. The CTOT index is defined as:

$$CTOT_{jt} = P^X_{jt}/P^M_{jt} = \prod_i (P_{it}/MUV_t)^{X^i_j} / \prod_i (P_{it}/MUV_t)^{M^i_j} \tag{1}$$

where $CTOT_{jt}$ is the CTOT index for country j in year t. P_{it} is a global price index for each of six commodity categories (food, fuels, agricultural raw materials, metals, gold, and beverages); X^i_j is the average export share of commodity i in GDP from 1970 to 2006; M^i_j is the corresponding average import share; and the commodity prices are deflated by a manufacturing unit value index (MUV) as in Deaton (1999) and Fernandez (2012). The fact that X^i_j and M^i_j are averaged over the sample years ensures that the CTOT index is invariant to changes in export and import volumes in response to conflict outcomes, thus isolating the effect of commodity price fluctuations. The growth rate of the CTOT index can be approximated as follows:

$$dlnCTOT_{jt} = dlnP^X_{jt} - dlnP^M_{jt} = \sum\nolimits_i X^i_j dln\tilde{P}_{it} - \sum\nolimits_i M^i_j dln\tilde{P}_{it} = \sum\nolimits_i NX^i_j dln\tilde{P}_{it} \tag{2}$$

where$\tilde{P}_{it} = \frac{P_{it}}{MUV_t}$ denotes the MUV deflated commodity price and NX^i_j is the average net export share of commodity i in GDP from 1970 to 2006. If we denote GDP in current U.S. dollars Y_t, deflate by the MUV, $\tilde{Y}_t = \frac{Y_{jt}}{MUV_t}$, abstract from non-commodity net-exports, and drop the country subscript, $d\tilde{Y}_t = d\left(\tilde{P}_{Ct}\tilde{C}_t + \tilde{P}_{It}\tilde{I}_t + \tilde{P}_{Gt}\tilde{G}_t + \sum_i NX^i_t\tilde{P}_{it}\right)$ in the usual notation. If commodity prices are uncorrelated with non-commodity spending, $d\tilde{Y}_t = d\sum_i NX^i_t\tilde{P}_{it}$. Dividing by last year's MUV-deflated GDP yields

$$\begin{aligned}\frac{d\tilde{Y}_t}{\tilde{Y}_{t-1}} &= \frac{d\sum_i NX^i_t\widetilde{P_{it}}}{\tilde{Y}_{t-1}} \approx \frac{\sum_i NX^i_{t-1}\tilde{P}_{it-1}}{\tilde{Y}_{t-1}} \frac{d\sum_i NX^i_t\widetilde{P_{it}}}{\sum_i NX^i_{t-1}\tilde{P}_{it-1}} \approx \frac{\sum_i NX^i_{t-1}\tilde{P}_{it-1}}{\tilde{Y}_{t-1}} \frac{d\sum_i NX^i_t\left(\widetilde{P_{it}} - \tilde{P}_{it-1}\right)}{\sum_i NX^i_{t-1}\tilde{P}_{it-1}} \\ &\approx \sum\nolimits_i NX^i_j d\ln\widetilde{P_{it}}\end{aligned} \tag{3}$$

if we assume that the share of net commodity exports in GDP is relatively constant over the sample years and the quantity response to price shocks is small in the impact year. Thus, the change in the MUV-deflated NX/GDP ratio and MUV-deflated economic growth should respond approximately one-for-one to the log CTOT index.[5]

Brückner and Ciccone (2010), Bazzi and Blattman's (2014), and Janus and Riera-Crichton (2015) note that changes in commodity prices may be serially correlated and have lagged conflict effects. In the estimation, therefore, we use the growth rate of the three-year moving average prices. For example, in the case of export prices,

$$\Delta P^X_{jt} = \ln\sum\nolimits_{s=t-2}^{t} P^X_{js}/3 - \ln\sum\nolimits_{s=t-3}^{t-1} P^X_{js}/3 \tag{4}$$

The appendix to Janus and Riera-Crichton (2015) shows that the growth rate of the three-year moving average price index approximately equals the average annual growth rate over the three years. We prefer the inclusion of the moving average index since we otherwise get a very large number of interaction terms as we decompose the index into export and import prices, positive and negative price shocks, and price shocks to extractive and non-extractive natural resources.

Table 1. Summary Statistics.

Variable	Obs	Mean	SD	Min	Max
Total Battle Deaths Low Estimate	900	1,202	2,908	14	37,000
Total Battle Deaths High Estimate	900	5,847	12,472	26	200,000
Total Battle Deaths Best Estimate	653	3,614	8,330	25	80000
Total Battle Deaths Imputed Best Est.	900	3,713	8,480	25	100,500
%Δ3-Year Commodity Terms of Trade (CTOT)	900	0.001	0.012	−0.071	0.111
%Δ3-Year Moving Average Export Prices	900	0.000	0.011	−0.061	0.079
%Δ3-Year Moving Average Import Prices	900	0.000	0.006	−0.030	0.027
Pre-war executive constraints	785	3.6	2.3	1	7
I(Pre-war executive constraints<median)	785	0.40	0.49	0	1

Regression Specification

Our regressions relate the natural logarithm of battle-related fatalities per country-year to the growth rate of 3-year-moving-average commodity export and import prices. In the main specification, we take the natural log of battle deaths since the variable is highly non-normally distributed and to reduce the probability that extreme observations affect the result. However, we later show that the findings are robust to excluding civil wars with a small number of battle deaths and estimating battle deaths per 1000 population. We control for conflict-episode fixed effects, region-specific year effects, and episode-specific quadratic time trends. We estimate robust standard errors clustered at the level of the conflict episode. A conflict episode is an unbroken series of conflict years. For example, if a country has battle-deaths observations in years (t) to (t + 2), (t + 4) to (t + 8), and (t + 10) to (t + 17), we record three different conflict episodes that last, respectively, 3, 5, and 8 years. Altogether, we estimate:

$$b_{ejt} = a + \beta^X \Delta P^X_{jt} + \gamma^X \Delta P^X_{jt} I\left(\Delta P^X_{jt} > 0\right) + \beta^M \Delta P^M_{jt} + \gamma^M \Delta P^M_{jt} I\left(\Delta P^M_{jt} > 0\right) + \mu_e + z_{rt} + \rho_e d_{et} + \phi_e d^2_{et} + \varepsilon_{ejt} \tag{5}$$

where b_{ejt} is the natural logarithm of the number of battle-related fatalities in episode e in country j in year t, ΔP^X_{jt} is the growth rate (or log change) in the three-year moving-average of commodity export prices; $\Delta P^X_{jt} I\left(\Delta P^X_{jt} \geq 0\right)$ is the growth rate multiplied by an indicator that equals one when the growth rate is positive and zero otherwise. If γ^X is statistically significant, then, positive and negative export price shocks have asymmetric effects. If γ^X is also positive, then, commodity export price stabilization can, potentially, reduce battle deaths. This is because the effect of a 1% decrease in export prices on battle deaths is $-0.01\beta^X>0$ if β^X is negative (Bazzi and Blattman 2014). The effect of a 1% *increase* in export prices is $0.01(\beta^X+\gamma^X)$. If prices go up or down with a 50% probability each, the effect of price volatility is $-0.01\beta^X/2 + 0.01(\beta^X+\gamma^X)/2 = 0.01\gamma^X\ /2 > 0$. ΔP^M_{jt} and $\Delta P^M_{jt} I\left(\Delta P^M_{jt} > 0\right)$ are, analogously, the growth rate in the three-year moving-average of commodity *import* prices and $\Delta P^X_{jt} I(\Delta CTOT_{jt} \geq 0)$ is the growth rate multiplied by an indicator that equals one when the growth rate is positive. If higher import prices decrease conflict, $\beta^M>0$, and if positive and negative import price shocks have asymmetric effects, γ^M should be statistically significant (we do not find evidence for this). μ_e is a conflict-episode-specific year effect. z_{rt} is a region-specific year effect for six different regions: developed countries (Western countries and Japan), Eastern Europe and the Former Soviet Union, Latin America and the Caribbean, Sub-Saharan Africa, Asia, and North Africa and the Middle East. $\left(\rho_e d_{et} + \phi_e d^2_{et}\right)$ is an episode-specific quadratic time-trend, where the duration variable-d_{et}records how long episode e has lasted in year t (e.g. the death toll might follow an arch-shaped pattern from the onset year). We only include conflicts with a known start date and the episode-specific time trends are defined relative to this start date and not relative to the beginning of the sample. Tables 1 and 2 display the summary statistics and sample countries.

Table 2. Sample Countries.

Afghanistan	Congo, Dem. Rep.	Haiti	Mexico	Philippines	Togo
Angola	Cote d`Ivoire	India	Moldova	Romania	Trinidad &Tobago
Argentina	Croatia	Indonesia	Morocco	Russia	Tunisia
Azerbaijan	Djibouti	Iran	Mozambique	Rwanda	Turkey
Bangladesh	Egypt	Iraq	Nepal	Saudi Arabia	Uganda
Bosnia & Herz	El Salvador	Israel	Nicaragua	Senegal	United Kingdom
Burkina Faso	Eritrea	Kenya	Niger	Sierra Leone	United States
Burundi	Ethiopia	Laos	Nigeria	Somalia	Venezuela
Cambodia	The Gambia	Lebanon	Oman	South Africa	Zimbabwe
Cameroon	Georgia	Lesotho	Pakistan	Spain	
Central Afr. Rep	Ghana	Liberia	Panama	Sri Lanka	
Chad	Guatemala	Malaysia	Papua New G	Sudan	
Chile	Guinea	Mali	Paraguay	Tajikistan	
Colombia	Guinea-Bissau	Mauritania	Peru	Thailand	

Table 3. Baseline Estimates.

	(1)	(2)	(3)	(4)	(5)	(6)
Sample	Full	Full	3+yrs into conflict	Annual BD data	No large Exporters	Full (quantile regression)
ΔP_X	−9.7	−40.7***	−42.0***	−70.5***	−40.6***	−34.3***
	[9.2]	[9.7]	[10.4]	[17.0]	[10.0]	[12.5]
ΔP_X *I(ΔP_X > 0)		66.0***	72.2***	70.6**	66.1***	84.1**
		[18.9]	[23.8]	[34.1]	[19.5]	[39.4]
ΔP_M	40.8**	38.2*	9.9	39.6	32.8	49
	[18.6]	[22.0]	[24.9]	[41.2]	[23.0]	[33.3]
ΔP_M *I(ΔP_M > 0)		41.6	84.1	30.8	46.8	37.3
		[48.0]	[68.0]	[60.5]	[52.5]	[107.2]
Observations	900	900	646	742	851	900
R-squared	0.74	0.75	0.73	0.78	0.74	0.67
# Conflict episodes	154	154	71	143	146	154
Conflict episode FE	Y	Y	Y	Y	Y	Y
Region×Year FE	Y	Y	Y	Y	Y	Y
Episode-specific quadratic time trend	Y	Y	Y	Y	Y	Y

Note: Cluster-robust standard errors in brackets. *, **, *** significant at 10%, 5%, 1%. Δ denotes the log change in the three-year moving average. The quantile regression is an unconditional regression for the 50th percentile (Firpo, Fortin, and Lemieux 2009; Borgen 2016).

Results

Table 3 , Column (1) indicates that higher import prices increase battle deaths (Janus and Riera-Crichton 2018). This is consistent with the idea that declines in real income and productivity in terms of importable goods (as households can afford fewer imports) increase conflict. In contrast, the effect of export prices is small and insignificant, which suggests that export prices do not monotonically increase battle deaths. In Column (2), we distinguish between positive and negative export and import price shocks and find that consistent with hypothesis H0a, positive and negative export price shocks appear to have asymmetric effects: γ^X, or the coefficient on ΔP_X*I(ΔP_X>0), is significantly positive. The sum of the coefficients in Table 3 is consistently (although, mostly insignificantly) positive. This suggests that, if anything, positive export price shocks might *increase* battle deaths (Dube and Vargas 2013; Berman et al. 2017).

In Column (3), we restrict the sample to observations three or more years into the conflict: we use the growth rate of the 3-year moving average export price index, and the effects of price shocks before and after the conflict onset might be different. Following Bazzi and Blattman's (2014), Column (4) restricts the sample to observations for which where Lacina and Gleditsch (2005) observe year-specific deaths instead of relying on interpolated or noisy estimates (Lacina 2009, 5). Column (5) excludes countries whose export share represents at least 5% of global exports in a commodity category. Column (6) estimates an unconditional (50^{th}) quantile regression, which is more robust to outliers. The unconditional regression identifies the marginal effect on the pre-regression quantile (Firpo, Fortin, and Lemieux 2009; Borgen 2016). The effects of positive export price shocks remain statistically different from the effects of negative price shocks.

In Table 4, we divide the sample into countries with weak and strong pre-conflict institutions. We measure institutional quality as the degree of constraints on the executive power in the Polity IV dataset and define an indicator that equals one when executive constraints immediately before the first conflict year is below the median (which is 3 on a discrete 1 to 7 scale). Table 4, Columns (1)-(2) imply that consistent with hypothesis H0b, negative and positive export price shocks have asymmetric effects when institutions are weak. Both types of shocks increase battle deaths. Additionally, positive shocks to import prices may increase battle deaths, although the p-value is only 0.15 in the quantile specification. In Column (3), we try to ensure that the coefficients are not identified from a small number of conflict episodes by re-estimating the quantile regression with year effects. The qualitative results remain similar and positive shocks to import prices now significantly increase battle deaths ($p = 0.00$). In contrast, Columns (4)-(6) show that commodity prices are insignificantly related to battle deaths when pre-conflict institutions are strong as opposed to weak.

In Table 5, we distinguish between the effects of extractive (fossil fuels, metals, and gold) and non-extractive (food, beverages, and agricultural raw materials) export price shocks. Columns (1)-(2) indicate that consistent with hypothesis H0c, positive and negative shocks to extractive commodity export prices have asymmetric effects when institutions are weak. In Columns (3) and (4), we again re-estimate Columns (1)-(2) with year effects. Columns (5)-(8) use two alternative measures of institutional quality. First, we define an indicator that equals one when pre-war executive constraints are either below *or equal to* the median of 3 on the discrete 1 to 7 scale. This definition increases the number of conflict episodes from 47 to 80. Second, we use an indicator that equals one when the economy's pre-conflict polity2 democracy score in the Polity IV dataset is equal to or below the median (−2 on a −10 to 10 scale). This produces 70 conflict episodes. The correlations between the

Table 4. Price effects under weak and strong institutions.

	(1)	(2)	(3)	(4)	(5)	(6)
Institutional Quality	Weak	Weak	Weak	Strong	Strong	Strong
Regression	LS	Quantile	Quantile (only year FE)	LS	Quantile	Quantile (only year FE)
ΔP_X	−78.2***	−111.6**	−39.5*	−28.2	54.2	−2.1
	[11.2]	[46.1]	[20.8]	[36.8]	[115.4]	[36.1]
ΔP_X *I($\Delta P_X > 0$)	131.7***	236.0***	138.8***	5.8	−126.3	−22.1
	[23.4]	[55.2]	[20.5]	[69.6]	[180.6]	[60.0]
ΔP_M	55.2*	61.3	60.2	25.4	−35.5	−22.9
	[29.8]	[124.1]	[59.9]	[60.6]	[143.1]	[99.5]
ΔP_M *I($\Delta P_M > 0$)	186.7	386	302.2*	−32.2	−189.6	−62.8
	[122.0]	[329.8]	[168.3]	[94.9]	[257.3]	[130.3]
Sum of first two terms	53.5	124.4	99.3	−22.4	−72.1	−24.2
p-value	0.03	0.00	0.00	0.73	0.68	0.59
Sum of last two terms	241.9	447.3	362.4	−6.8	−225.1	−85.7
p-value	0.06	0.15	0.00	0.94	0.43	0.46
Observations	313	313	313	472	472	472
R-squared	0.84	0.8	0.73	0.81	0.71	0.61
# Conflict episodes	47	47	47	89	89	89
Conflict episode FE	Y	Y	Y	Y	Y	Y
Episode-specific qtt	Y	Y	Y	Y	Y	Y
Region×Year FE	Y	Y	N	Y	Y	N
Year FE	Y	Y	Y	Y	Y	Y

Note: Cluster-robust standard errors in brackets. *, **, *** significant at 10%, 5%, 1%. Δ denotes the log change in the three-year moving average. The quantile regression is an unconditional regression for the 50^{th} percentile (Firpo, Fortin, and Lemieux 2009; Borgen 2016).

Table 5. Price effects of extractive and nonextractive exports in the weak institutions sample.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Weak Institutions proxy	Pre-conflict ex const < median		Pre-conflict ex const < median		Pre-conflict ex const ≤ median		Pre-conflict polity2 ≤ median	
Time trends include	Region×Year Effects		Year Effects		Region×Year Effects		Region×Year Effects	
Regression	LS	Quantile	LS	Quantile	LS	Quantile	LS	Quantile
ΔP_{XE}	−71.5***	−89.6**	−27.0***	−28.7*	−70.6***	−45.9**	−71.0***	−48.8***
	[10.7]	[39.3]	[5.8]	[15.8]	[10.2]	[17.4]	[10.1]	[17.2]
ΔP_{XE} *I(ΔP_{XE} > 0)	108.0***	168.4***	53.7***	90.8***	107.4***	153.1***	107.5***	154.3***
	[14.5]	[39.2]	[8.1]	[18.7]	[14.3]	[22.1]	[14.2]	[22.6]
ΔP_{XN}	29.1	−284.5	117.6	−59	60.3	83.1	36	−60.2
	[97.5]	[332.1]	[73.0]	[234.2]	[47.3]	[145.4]	[48.5]	[90.3]
ΔP_{XN} *I(ΔP_{XN} > 0)	−400.9*	29.5	−97.3	335.3	−43.6	141.8	−9.4	280.8
	[209.8]	[565.5]	[174.3]	[393.0]	[145.9]	[366.9]	[138.8]	[331.5]
ΔP_M	74.9***	101.6	66.7***	143.7***	56.3	77.5	56.3	63
	[23.9]	[90.1]	[15.2]	[42.9]	[35.6]	[73.1]	[36.2]	[75.1]
Sum of first two terms	36.5	78.8	26.7	62.1	36.8	107.2	36.5	105.5
p-value	0.01	0.00	0.00	0.00	0.01	0.00	0.01	0.00
Observations	313	313	313	313	479	479	438	438
R-squared	0.84	0.8	0.75	0.72	0.82	0.74	0.82	0.71
# Conflict episodes	47	47	47	47	80	80	70	70
Conflict episode FE	Y	Y	Y	Y	Y	Y	Y	Y
Episode-specific qtt	Y	Y	Y	Y	Y	Y	Y	Y
Region×Year FE	Y	Y	N	N	Y	Y	Y	Y
Year FE	N	N	Y	Y	N	N	N	N

Note: Cluster-robust standard errors in brackets. *, **, *** significant at 10%, 5%, 1%. Δ denotes the log change in the three-year moving average. The quantile regression is an unconditional regression for the 50th percentile (Firpo, Fortin, and Lemieux 2009; Borgen 2016). XE and XN denote extractive (fuels, metals, and gold) and non-extractive (food, beverages, and agricultural raw materials) exports.

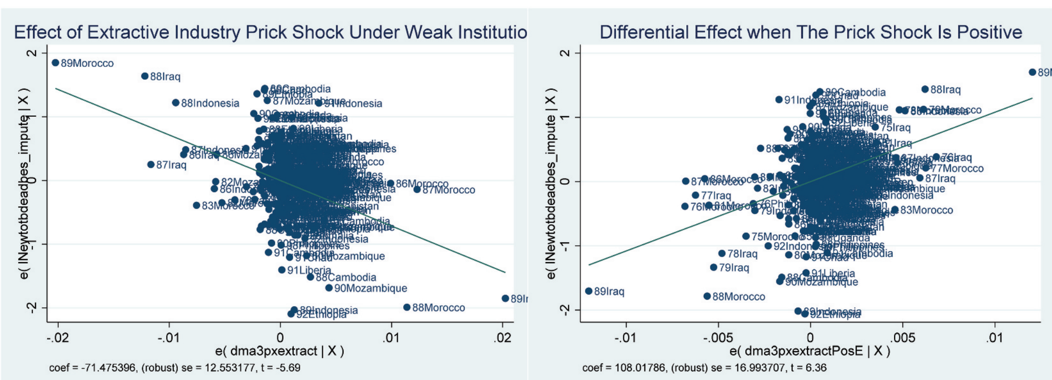

Figure 1. Effects of negative and positive extractive price shocks under weak institutions. *Note*: The figure depicts the added-variable plots from the regression in Table 5, Column (1).

original and the two alternative institutional quality indicators range from 0.65 to 0.88. Figure 1 depicts the added-variable plots for the effects of extractive price shocks and the interaction term when the price shock is positive in Table 5, Column (1).

In Table 6, we decompose the extractive price shock into price shocks to fossil fuels, metals, and gold. As we disaggregate the price shocks, the likelihood that a relatively small number of conflict observations explain the individual coefficients increases. However, fossil fuels play an outsized role in the global economy and for many individual countries and governments, so the effects of extractive price shocks may be heterogenous or mainly explained by fuel prices. Columns (1)-(2)

Table 6. Price effects of disaggregated extractive export price shocks in the weak institutions sample.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Controlling for region-specific year effects				Controlling for year effects			
Weak Institutions measure	Pre-conflict executive constraints < median		Pre-conflict executive constraints ≤ median		Pre-conflict executive constraints < median		Pre-conflict executive constraints ≤ median	
Regression	LS	Quantile	LS	Quantile	LS	Quantile	LS	Quantile
ΔP_{XFUEL}	−68.7***	−71.5**	−65.2***	−34.7***	−22.2***	−20.6	−26.6***	−29.5***
	[7.3]	[28.0]	[6.8]	[12.6]	[5.0]	[12.7]	[4.6]	[6.5]
ΔP_{XFUEL}*I(ΔP_{XF} > 0)	109.4***	150.2***	107.1***	150.1***	49.3***	80.2***	53.8***	80.2***
	[14.6]	[32.6]	[14.4]	[21.1]	[9.5]	[18.6]	[8.0]	[11.0]
ΔP_{XMETAL}	−59.8	−380.6***	−70.3**	−105.0**	−44.4	−291.1***	−36.8**	−64.2*
	[55.9]	[93.1]	[33.5]	[41.6]	[27.6]	[50.2]	[14.1]	[37.1]
ΔP_{XMETAL}*I(ΔP_{XM} > 0)	−79.9	554.2**	−55.8	17.1	−39.3	490.4***	−49.3**	15.2
	[73.4]	[240.0]	[48.3]	[101.0]	[39.0]	[92.3]	[24.1]	[163.4]
ΔP_{XGOLD}	−3,352.2***	−4,426.00	−2,147.00	811.3	−1,955.10	−2,318.10	−1,082.50	2,468.20
	[1,110.1]	[3,692.8]	[1,401.9]	[2,099.7]	[1,300.1]	[2,664.7]	[1,103.1]	[1,558.7]
ΔP_{XGOLD}*I(ΔP_{XG} > 0)	4,697.4***	5,342.40	3,088.7*	−749.1	2,467.30	2,224.10	1,367.30	−3,115.30
	[1,405.9]	[4,120.6]	[1,850.2]	[2,429.0]	[1,993.1]	[3,610.8]	[1,665.4]	[1,893.3]
ΔP_{XN}	−85.2	−300.8	39.9	98.8	82.5	6.6	54.7**	22.7
	[85.1]	[274.4]	[49.4]	[121.9]	[87.9]	[252.5]	[23.5]	[53.3]
ΔP_M	81.3***	98.9	60.6*	71	68.2***	128.3***	59.2***	87.1***
	[22.4]	[94.7]	[33.7]	[71.0]	[12.6]	[37.5]	[12.8]	[30.9]
Observations	313	313	479	479	313	313	479	479
R-squared	0.85	0.82	0.83	0.74	0.76	0.73	0.74	0.64
# Conflict episodes	47	47	80	80	47	47	80	80
Conflict episode FE	Y	Y	Y	Y	Y	Y	Y	Y
Episode-specific q tt	Y	Y	Y	Y	Y	Y	Y	Y
Region×Year FE	Y	Y	Y	Y	N	N	N	N
Year FE	N	N	N	N	Y	Y	Y	Y

Note: Cluster-robust standard errors in brackets. *, **, *** significant at 10%, 5%, 1%. Δ denotes the log change in the three-year moving average. The quantile regression is an unconditional regression for the 50th percentile (Firpo, Fortin, and Lemieux 2009; Borgen 2016). XN denotes non-extractive (food, beverages, and agricultural raw materials) exports.

decompose Table 5, Column (1). Columns (3)-(4) re-estimate the regression when, alternatively, we define weak institutions as the existence of executive constraints that are below *or equal* to the median. Columns (5)-(8) reproduce Columns (1)-(4) when we only control for year effects to avoid overfitting the model. Table 6 shows that the asymmetric effects of positive and negative extractive price shocks are only consistently positive and significant for fossil fuels. However, we would hesitate the conflict that gold and metal price shocks do not have asymmetric effects. Instead, our study might be underpowered to detect these asymmetric effects: partly, the global volatility of energy prices tends to exceed the volatility of other commodity prices (Regnier 2007) and partly the average export share of fossil fuels greatly exceeds the export shares of metals and gold, which increases the identifying variation.

Robustness Checks

Other Dependent Variables

In the supplementary appendix, Table A1, Columns (1)-(2) replace the dependent variable in Table 4, Column (1) with the log of the highest and lowest estimated death tolls in Lacina and Gleditsch (2005). In Columns (3)-(4), we consider whether the fact that we log the dependent variable explains or disproportionally influences the results: when we take the log, the same unit change in log battle deaths can in principle be achieved if battle deaths increase from 25 to 50 or from 1000 to 2000. We use two methods to address this concern. First, we restrict the sample to observations where the battle-related death toll in the previous conflict year exceeds 500. The

500 threshold balances our desire to retain a moderately large sample (which is reduced by a third), number of conflict episodes (which is reduced by 45%), and degrees of freedom after controlling for episode-specific fixed effects, quadratic episode-specific time trends, and region-specific year effects. Although the least-squares estimate for ΔP_{XE}*I(ΔP_{XE}>0) in Column (3) is marginally insignificant ($p = 0.14$), it retains a large positive magnitude. Additionally, Column (4) shows that the estimate regains significance in the quantile regression ($p = 0.04$), which is considered more robust to outliers. Second, we estimate total (non-logged) battle death per 1000 individuals in the population.[6] Since this variable is highly non-normally distributed, we drop the highest 5% of the observations. Columns (5) and (6) report the least-squares and quantile estimates. The estimate for ΔP_{XE}*I(ΔP_{XE}>0) remains significant. In Column (7) we estimate the log of battle deaths per 1000 population as opposed to log total battle deaths. The idea is that larger populations might suffer more battle deaths because the population is larger. Finally, Columns (8) and (9) first-differences the regression for log battle deaths and log deaths per 1000 population to control for time-series persistence. The results for ΔP_{XE}*I(ΔP_{XE}>0) remain significant (positive and negative price shocks still have asymmetric effects), although the effect of positive shocks loses significance ($p = 0.18$ in both regressions).

Omitting Individual Regions

Table A2 presents the results of omitting individual regions. When the North African and Middle Eastern region is omitted in Column (6), the standard error on the extractive industry price shock increases, which reflects that we drop Middle Eastern fuel exporters. This decreases the identifying variation and decreases the signal-to-noise ratio. Nonetheless, the coefficient on the interaction term, ΔP_{XE}*I(ΔP_{XE}>0), remains significant ($p = 0.09$). Thus, positive and negative price shocks still have asymmetric effects. In Column (7), moreover, we show that if we replace the region-specific year effects with year effects (representing time-varying global rather than time-varying regional shocks), the estimate regains high significance ($p < 0.01$) without markedly decreasing the model fit (R-squared decreases from 0.85 to 0.82). We also re-estimated the model without North Africa and The Middle East as a quantile regression. The statistical significance level on ΔP_{XE}*I(ΔP_{XE}>0) was then $p = 0.16$ with region-specific year effects and $p = 0.03$ with year effects.

Control Variables

The fact that we control for episode-fixed effects (a fixed effect for each unbroken conflict episode within a country) implies that we control for time-invariant conflict determinants, such as colonial history, pre-conflict GDP per capita, and other pre-conflict economic, political, geographic, or demographic variables. However, we have not controlled for time-varying determinants of battle deaths. Unfortunately, many domestic time-varying factors, such as income per capita and political exclusion during the conflict may be endogenous to battle deaths or represent the causal mechanism linking commodity prices to battle deaths. Therefore, we only try to control for foreign conflict participation (which may or may not be endogenous) and for temperature and rainfall shocks (Burke, Hsiang, and Miguel 2015). We use the fact that Lacina and Gleditsch (2005) distinguish between internal and internationalized internal armed conflict and define an indicator for foreign participation that equals one if at least one of the ongoing civil conflicts within a country-year observation is internationalized. This is true for 15% of the observations.[7] In Table A3, Columns (1) and (2) show that the least-squares and quantile estimates for ΔP_{XE}*I(ΔP_{XE}>0) remain robust.

In Table A3, Columns (3)-(4), we use population-weighted rainfall and temperature data from Bazzi and Blattman's (2014), who source the data from Dell, Jones, and Olken (2012). Following Burke, Hsiang, and Miguel (2015), we standardize the temperature and rainfall deviations by subtracting the country-specific mean for country j (denoted μ_j) and dividing by the standard deviation (σ_j) in Bazzi and Blattman's (2014) dataset, which covers 118 countries from 1957-2007.

For example, the standardized temperature deviation (z-score) for country j in year t is $z\_temp_{jt} = \left(temp_{jt} - \mu_j\right)/\sigma_j$. We control separately for the effects of positive and negative deviations by defining $z^{+}_{jt\_temp} = \max\left(z_{temp_{jt}}, 0\right)$ and $z^{-}_{jt\_temp} = \min\left(z_{temp_{jt}}, 0\right)$. This is equivalent to controlling for z_temp_{jt} and $(z\_temp_{jt}) * I\left(z_{temp_{jt}} > 0\right)$ except that testing for asymmetric effects implies testing whether the coefficients on $z^{+}_{jt\_temp}$ and $z^{-}_{jt\_temp}$ are equal. Following Burke, Hsiang, and Miguel (2015), we control for contemporary and lagged positive and negative temperature and rainfall deviations. Finally, Columns (5) and (6) control jointly for foreign participation, rainfall, and temperature deviations. The estimates for ΔP_{XE}*I(ΔP_{XE}>0) remain significantly positive.

In Table A4, we control for non-linear, non-institutional effects to study whether price shocks have heterogenous effects due to reasons other than institutional quality. As a benchmark, Columns (1) and (2) present the least-squares and quantile estimates when we estimate the full sample instead of dividing it into countries with weak and strong institutions. The fact that ΔP_{XE}+ΔP_{XE}*I(Weak Institutions) is significantly negative implies that more negative extractive price shocks increase battle deaths in countries with weak institutions. The fact that ΔP_{XE}*I(ΔP_{XE}>0)+ΔP_{XE}*I(ΔP_{XE}>0)*I (Weak Institutions) is significantly positive implies that positive and negative extractive price shocks have asymmetric effects in countries with weak institutions. In the remaining columns, we control for non-linear interactions to test whether the apparent effects of weak institutions are better explained by poverty or ethnic divisions. Thus, we construct indicators that equal one when pre-conflict PPP-adjusted income per capita (Feenstra, Inklaar, and Timmer 2015) is below the median and when ethnic fractionalization and polarization from Fearon (2003) (and calculated following Janus and Riera-Crichton (2015)) exceed their medians. Then, we add interactions between positive and negative extractive price shocks and the relevant indicators. The sign, magnitude, and significance of ΔP_{XE}+ΔP_{XE}*I(Weak Institutions) and ΔP_{XE}*I(ΔP_{XE}>0)+ΔP_{XE}*I(ΔP_{XE}>0)*I(Weak Institutions) generally remains robust. When we control for high ethnic polarization in Columns (7) and (8), one of the sums loses significance. Nonetheless, the signs of the individual coefficients and the coefficient sums remain similar to, respectively, Columns (1) and (2). In contrast, the polarization controls change signs across regressions (7) and (8) and are individually and jointly insignificant.

Conclusion

This paper hypothesized that positive and negative commodity price shocks may have asymmetric effects on conflict. First, we developed a theoretical framework that explained or summarized (Moshiri 2015) five possible causes of asymmetry and why weak institutions and extractive-industry dependence might increase the asymmetry. Second, we derived and tested three asymmetric-effects hypotheses in a dataset of battle deaths during ongoing civil wars from 1970-2008. Consistent with the hypotheses, negative price shocks increase battle deaths and positive shocks also have a positive effect. The positive effect is concentrated and becomes statistically significant in countries that have weak pre-conflict institutions (defined as unconstrained or undemocratic governments) and experience a positive price shock to fossil fuels. We conclude that a combination of institutional reforms, economic diversification, and price stabilization (Van der Ploeg and Poelhekke 2009; Frankel 2017) might reduce battle deaths in fuel-dependent countries with weak institutions.

Notes

1. Another literature studies the effect of natural resources on democracy, corruption, and other institutional quality aspects (Vicente 2010; Ross 2012; Waldner and Smith 2015; Cassidy 2019). However, in this paper, we assume that institutional quality is pre-determined and measure institutional quality before the conflict.
2. Several observations suggest that it is important to treat the onset (i.e. peace to war transitions) and the intensity of conflict conditional on the onset as separate research questions. For example, Bazzi and Blattman's (2014, abstract) conclude that 'Conflict onset and continuation follow different processes.' Intuitively, if civil wars cause

institutional breakdowns, they can change the way the economy responds to economic shocks. For example, during the civil wars in Afghanistan, Colombia, Iraq, Libya, Somalia, Sri Lanka, and Syria over the last 60 years, non-armed non-state groups carved out large autonomous regions or 'parallel states' where they collected revenues through taxation, extortion, kidnapping, smuggling, and natural resource extraction and exports. The revenues were often used to finance the war (Keen 2000; Rubin 2000; Le Billon 2001; Bannon and Collier 2003).

3. Although the conflict literature has increasingly moved to micro-level data, macroeconomic conflict studies can also have advantages (Balcells and Stanton 2021). First, it is difficult to get subnational institutional quality data and institutions might vary less within countries. Second, it is difficult to construct comprehensive commodity export and import price data for subnational units and commodity prices are at least unconditionally correlated (Erten and Ocampo 2013). Third, subnational analysis might increase the probability of general equilibrium and strategic effects, such as population displacement or combatants shifting from neighboring units. Related, our definition of civil war requires that the central government is a party, which might make the country a 'natural' unit of analysis.
4. Terms of trade growth should normally increase welfare as each export good can be traded for more import goods (Mendoza 1995). Lederman and Porto (2016) discuss the effect of commodity prices on household welfare based on survey data from Africa and Latin America as well as a review of the literature. They conclude that households spend large budget fractions on commodities, they often depend on commodities to earn income, and international price changes pass through to households, suggesting households are exposed to both import and export prices.
5. The reason why the United States is included in the sample in Table 2 is that Lacina and Gleditsch (2005) count the death toll in Afghanistan after the terrorist attacks on U.S. soil on 9/11/2001 as an internationalized internal conflict.
6. The population data is from the Penn Tables Version 9.1 (Feenstra, Inklaar, and Timmer 2015), supplemented with World Bank W.D.I. and U.N. Population Division.
7. Ideally, we would have liked to construct a randomized measure of intervention (since the decision to intervene may be endogenous) and obtain information about the resources and policy objectives of the intervener.

Disclosure statement

No potential conflict of interest was reported by the authors.

Data Availability Statement

The dataset and replication code for this paper is located at https://doi.org/10.6084/m9.figshare.21191515.v1

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