FOOD PRICE SPIKES AND SOCIAL UNREST IN AFRICA[1]

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INTRODUCTION

Observed food riots throughout much of Africa and the Middle East in the wake of spikes in international food commodity prices have fueled a renewed academic and popular interest in the connection between food prices and social unrest, including widespread speculation and debate about the role of food prices in the Arab Spring. (Ciezadlo 2011; BBC 2008; Harsch 2008; BBC 2011; IRIN 2011; Aljazeera 2012) Dramatic weather events, diversion of crops for fuel production, and continuing volatility in international grain markets have led many to postulate that food riots will become more frequent, pervasive, and disruptive in the future. One group of researchers has published a report that has been picked up in the popular media that suggests that if international commodity prices continue to rise the world will become engulfed in riots by August of 2013. (Lagi, Bertrand, and Bar-Yam 2011; Brandon 2011; The Physics arXiv Blog 2011)

The proposition that rising food prices can and often do lead directly to food riots is widely accepted and not particularly controversial in either academic studies of political stability or

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political and economic policy calculations. It is, nonetheless, rather axiomatic. Prognostications based on such a simple correlation are, however, dangerous for a couple of reasons. First, it ignores the undoubtedly complex causal mechanisms that might link international food prices to social unrest. The most obvious connection is that rising international commodity prices drive up domestic consumer prices, which leads people to take to the streets under increased economic pressure, but the connection between international prices and local prices is not always clear. Second, such predictions disregard the endogenous nature of the relationship between food prices and unrest. Not only can rising food prices cause unrest, but unrest can also drive up food prices. Furthermore, both rising prices and unrest can be the result of a third factor, such as rising fuel prices or political corruption.

In this paper I seek to unravel the endogenous relationship between food prices and social unrest in Africa and provide some insight into potential causal mechanisms. I do so by introducing a number of theoretical and methodological innovations. From a theoretical perspective, I first ask the previously unaddressed question of whether rising food prices are a contributing causal factor not just of food riots but also of social unrest more broadly defined, such as labor strikes, student demonstrations, communal conflict, and other riots regardless of cited grievance. Secondly, I use domestic consumer food price indices as the independent variable because they better represent the prices that people actually pay for food and thereby provides some insight into the economic causal mechanisms.

Methodologically, I analyze the question at a unit of analysis and coverage not previously used, 40 countries in individual months over a period of 21 years. And finally, I employ an instrumental variable analysis to address the endogenous relationship between food prices and social unrest. In so doing I test two potential instrumental variables: international commodity prices and local rainfall deviations.

In the end, I find that a sudden increase in consumer food prices in a given month does contribute significantly to an increase in the probability of unrest in that month. Because of the instrumental variable specification I argue that this is strong evidence of a causal relationship. I also find that rainfall scarcity functions much better as an instrumental variable for changes in consumer food prices than do international commodity prices. Although more research is necessary to determine why people choose particular protest methods and targets, I argue that this is evidence that the causal mechanism at work is a response to economic pressure from rising prices, regardless of perceived injustices in the economy or food system. This provides some evidence, although not conclusive that the mechanism that links rainfall to unrest is through food prices but that any observed correlation between international commodity prices and food riots is not through local food prices and, therefore, needs another causal explanation.

This paper will proceed as follows. I will first provide an overview of the literature on food prices and food riots throughout a few periods in history. I will then provide a theoretical framework for the analysis. The next section will give a description of the data and the empirical strategy. This is followed by a presentation of the results. Finally, I will conclude with some thoughts about policy implications and directions for future research.

LITERATURE

Food riots in history

Historians, political scientists, and sociologists have all documented and studied the historical role of food and food prices on revolution, rioting, and unrest. Much has been written about the bread riots in eighteenth and nineteenth century England and France. Rudé (1964) thoroughly documents the significant role of food riots leading up to political upheavals in both countries. In France, bad harvests and lean years of 1725, 1740, 1749, 1768, 1775 and 1785 were accompanied by major popular disturbances. Similarly, failed harvests in England in every decade of the last half of the 18th century led to corresponding food price peaks and attendant outbursts of food riots in both the countryside and urban centers.

It would be a mistake, however, to think that these riots, or protests, were all of the same type, carried out by crowds of similar composition, or with common targets of discontent. Food riots took on many different forms with many different objectives including “simple looting of warehouses, attacking the homes of merchants, or stopping grain vessels bound for other parts." (Rudé 1964, 38)

Charles Tilly (1978) identifies three different types of food riots in England: the *retributive action*, an act committed by crowds against people believed to be hoarding or profiteering; the *blockage*, the prevention of food from leaving the area of production; and, the *price riot*, the seizure of food to be sold publicly at a proper price as set by the people. (Rudé 1964, 185) (See also Stevenson 1974.) Louise Tilly (1971) categorizes French protests into three types: *market riots*; *entrave*, "the rural form of grain riot, in which wagons or barges loaded with grain were forcibly prevented from leaving a locality"; and *taxation popularie*, a term referring to "the imposition of an unofficial price control by collective action." (Rudé 1964, 23)

Rudé (1964) also makes the point that the food riot was a more common form of social unrest than the strike amongst many industrial workers. Later, however, "these workers, although more prone in times of hardship to seek redress from grain factors and millers than from their own employers, were also on occasion involved in purely industrial disputes." (Rudé 1964, 66) Stevenson (1974) makes the same point and contends that these food riots eventually gave way to other forms of protest.

This illustrates that ascertaining the real driver of unrest or predicting the target can be a difficult task. Grievances about the cost of food can often be commingled with other grievances, economic and political, and the occurrence of a particular type of unrest is often determined more by a viable target and perceived culpability than the source of hardship. Goldstone (1982) writes: "Two grievances stand out as the chief cause of revolutionary urban tumults: the cost of food and the availability of employment." These causes are not mutually exclusive; rather, they are often intertwined with other grievances. How aggrieved groups choose the targets of their ire is a question that is largely outside the scope of this paper but is undoubtedly an important theoretical question with obvious practical import.

Finally, although individual outbreaks of unrest coincided with bad harvests and price spikes, the broader trend of food riots in eighteenth century was not a consequence of food shortage; they emerged at a time when per capita food supply was probably on the rise. Rather, they were attributable to three trends that accompanied the rise of the nation state and the increasing salience of the market in food systems. First, the proportion of European households that no longer produced a significant portion their own food, but rather relied on the market for survival greatly expanded. Second, food production became increasingly commercialized, which included an increasing prioritization of national markets over local needs. Third, “the extensive previously existing controls over the distribution of food, which gave the local population a prior claim over food produced and sold in a locality, and bound the local authorities to provide for the subsistence of the local poor” were dismantled. (C. Tilly 1978, 186)

Food riots and the modern food system

Food riots largely disappeared from the world stage from the mid-nineteenth century through the 1970s. (Rudé 1964; C. Tilly 1978; Walton and Seddon 1994) Beginning in the late 1970s, however, widespread unrest, including food riots, erupted in countries subjected to austerity programs imposed by international financial institutions in the wake of the debt crisis of the early 1970s. (Walton and Seddon 1994)

In the context of changes to food systems, the period of “structural adjustment” was the historical analogue of the late eighteenth century. Walton and Seddon (1994) document unrest throughout the developing world and provide detailed case studies of countries in Latin America, Asia, the Middle East and North Africa, and sub-Saharan Africa (Zambia, Côte d’Ivoire, and Zaire). They maintain “that modern food riots in the developing nations are generated by processes analogous to economic liberalization policies that produced classical food riots, but today’s transformation is taking place at the international level. Neo-liberalism simultaneously affects all Third World countries in much the same fashion as laissez-faire policies within nations once affected particular towns and regions, although the two processes are distinct in other ways.” (Walton and Seddon 1994, 24)

In order to provide an analytical lens to their case studies, they identify four general theories “to explain the origins, timing, and changing forms of food riots: rational response, moral economy, community, and state and market." (Walton and Seddon 1994, 30) *Rational response* theory, as developed by Bohstedt (1983), argues that regardless of the cause of rising prices or inequitable distribution disadvantaged groups respond rationally and proportionately to hardship by seizing food directly or appealing to authorities for relief. *Moral economy* theory, as developed by Thompson (1971), argues that protesters react against changes in existing food systems that are perceived to be unjust or illegitimate. Rather than focus on social classes or groups, proponents of the *community* theory emphasize traditional communities and the disruption of traditional ways of life as the mobilizing incentives for rioting. The *state and market* theory views as disruptive structural changes in the food system brought on by political decisions regarding state policies or regulation.

Ultimately, Walton and Seddon (1994) conclude that each of these theories provides an important element in explaining the motivation behind the food riots that were part and parcel of the *austerity protests* of the 1970s and 1980s, a broader expression of dissatisfaction with structural changes occurring both at the national and international levels.

> Modern bread riots, like their eighteenth-century predecessors, arise at the intersection of state and economic reorganization, now determined primarily at the international level.
> …
> The food riot as a means of popular protest is a common, perhaps even universal, feature of market societies – less a vestige of political-industrial evolution than a strategy of empowerment in which poor and dispossessed groups assert their claims to social justice. In the modern system of states and international economic integration, the explosive point of popular protest has moved, with most of the world's population, to the cities where the processes of global accumulation, national development, and popular justice intersect. (Walton and Seddon 1994, 53–54)

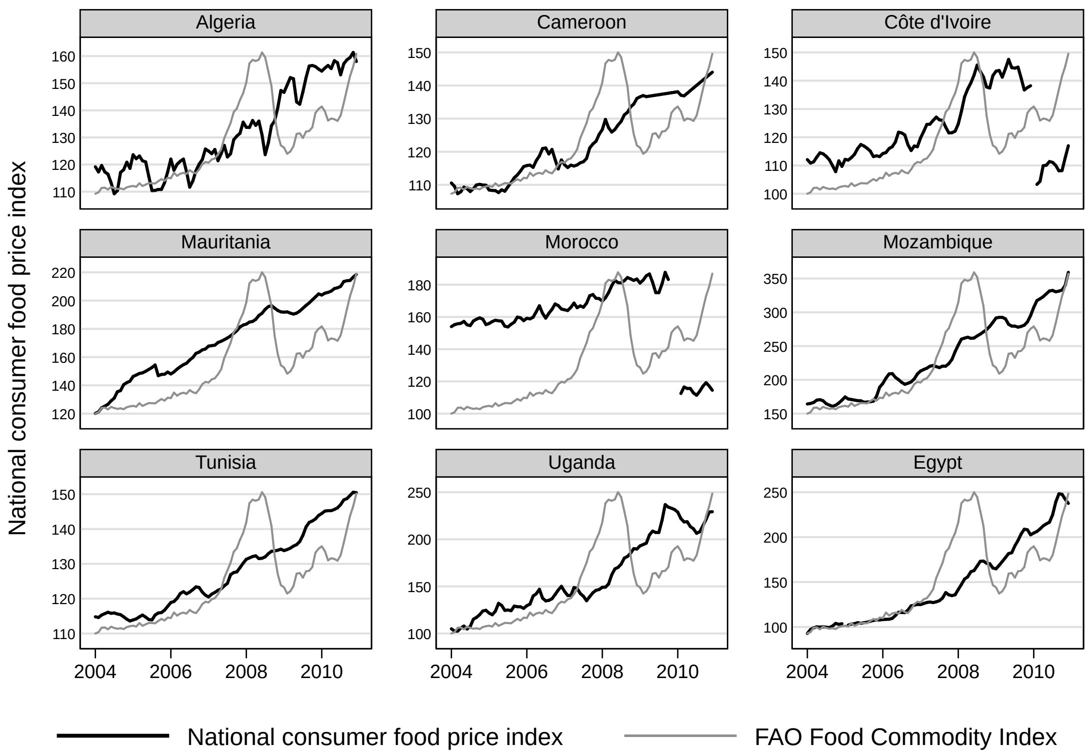
The more recent rash of food riots in the wake of the 2007–08 spike in international commodity prices has been repeatedly referred to as a new "food crisis." But, like both the eighteenth century and the 1980s, it is not a crisis of production. Record grain harvests were reported in 2007. The problem is one of *access* to food. (Holt-Giménez 2008)

In the year preceding the Arab Spring, Bush (2010) provides an insightful examination of the unrest during the preceeding two years. Like Walton and Seddon, he argues “the price increases also provided a catalyst for political mobilization and an opportunity to voice out dissent about a much broader range of concerns.” (Bush 2010, 119) He argues at least as important as the economic pressure of increasing prices was the perception of injustice in the global food system. He claims that “rioters knew why food prices were high and who benefited from food inflation. Rioters knew too why governments had to be forced to mitigate the social costs of food inflation, why and how authoritarian regimes appeased transnational food companies, and how national food strategies impoverished food producers: low farm-gate prices were well-tested mechanisms to extract surplus for largely urban-based development.” (Bush 2010, 121) He also recognizes that food riots were in some cases commingled with labor strikes. Rising food prices in Egypt led to a strike in April 2008 in which workers demanded higher wages to cover the increasing cost of food. (Bush 2010)

An emerging large-N literature has developed in the wake of the international food commodity spikes of 2007–08 and 2010-11 that seeks to identify a statistical relationship between food prices and political and social unrest. It is this literature to which the current paper most directly contirbutes, although I hope that it provides larger insights and a direction for future research into the theoretical motivations of protesters.

Relying largely on the theoretical framework of Acemoglu and Robinson (2001) that transitory economic shocks open a “democratic window of change,” Arezki and Brückner (2011) investigate the connection between food prices, political institutions, and political stability. They develop an index of annual national food prices for 120 countries from 1970 to 2007 and relate it to data on political institutions from the Polity IV project, civil conflict data from PRIO/Uppsalla as well as data on riots and demonstrations. They find political institutions in low-income countries deteriorate significantly during times of international food price increases largely because food prices increase “significantly increased the likelihood of civil conflict and other forms of civil strife, such as anti-government demonstrations and riots.” (Arezki and Brückner 2011, 11) The same was not true for high-income countries. They explain that although increases in international food prices have the macroeconomic effect of increasing real GDP in food-exporting countries, they also decrease consumption and increase income inequality in low-income countries.

Figure 1: National Consumer Food Indices for Countries with Observed Food Riots



Côte d'Ivoire series changed in January 2010; Morocco series changed in January 2010

Lagi et al. (2011) identify a correlation between international food prices, as measured by the FAO Food Index, and media reports of food riots. They argue that food security is the political responsibility of the state and that failure to provide this security by maintaining stable food prices leads to a break down of state authority and political order. “Conditions of widespread threat to security are particularly present when food is inaccessible to the population at large. In this case, the underlying reason for support of the system is eliminated, and at the same time there is “nothing to lose,” i.e. even the threat of death does not deter actions that are taken in opposition to the political order. Any incident then triggers death-defying protests and other actions that disrupt the existing order.” (Lagi, Bertrand, and Bar-Yam 2011, 3) They present a compelling but potentially misleading graph of the FAO Food Index with counts of food riots in individual countries. Upon closer examination, local consumer food indices in the individual countries included on the graph do not, in most cases, follow the same trend line as the FAO Index. Figure 1 illustrates this point using all available national consumer price indices from African countries included on Figure 1 of Lagi et al. (2011). Only in Côte d'Ivoire and Cameroon can any tracking be identified and this is delayed. The increases observed in international markets had not yet affected consumers in the countries where riots occurred, at least not prior to the rioting. This calls into question the proposed causal mechanism regardless of the correlation.

Berazneva and Lee (2011) conduct an analytical comparison of the countries in Africa that did and did not experience food riots during the 2007–08 price spike. They find higher levels of poverty and political repression are associated with a higher likelihood of riots. This is certainly an interesting finding and may contribute to understanding of the connection during 2007–08, but because this analysis is limited to a particular episode of global economic crisis its generalizability is questionable and it does not explain any potential boarder causal connection between domestic food prices and rioting.

Bellemare (Bellemare 2012) examines the link between monthly data on international food prices from the UN Food and Agricultural Organization (FAO) and media reports on food riots. He makes what amounts to a rational response theory argument for the causal connection. He writes: "In developing countries … substantial food price increases can lead to welfare losses that are so large as to threaten those poor households' very subsistence. But desperate times call for desperate measures, which means that the people in those households can take to the street if their livelihoods are threatened." (Bellemare 2012, 2) He recognizes that food riots can also lead prices increases and uses an instrumental variable approach to eliminate any potential bias from the endogeneity between food prices and riots. His chosen instrument is the occurrence of natural disasters in the International Disaster Database (EM-DAT).[3] He finds that increased food prices lead to increased political instability but that increased price volatility is associated with a

[3] Centre for Research on the Epidemiology of Disasters. "International Disaster Database (EM-DAT)," 2009. http://www.emdat.be/database.

decrease in political unrest. This paper investigates the correlation between food prices and food riots at the more appropriate monthly temporal unit and admirably attempts to address endogeneity, although with a questionable instrument which ignores that a "disaster" is not a purely exogenous product of natural events – a drought is not necessarily a famine and a famine is not necessarily a product of drought. Of concern, however, is that it still assumes that changes in international commodity prices are reflective of the prices that people pay domestically, which, as has been demonstrated, is not necessarily true. Additionally, Bellemare admits that his results are potentially driven by events in 2007–08 and 2010–11, which could suggest a different causal mechanism. Although the instrumental approach should eliminate bias from endogeneity, at the time both phenomena, rising prices and rioting, were potentially driven by other factors such as the global economic crisis or the Arab Spring.

Finally, all of these papers, with the exception of Arezki and Brückner (2011) are concerned exclusively with food riots. Historically, unrest over food prices has been commingled with other issues and grievances. It is not realistic to assume that food riots are solely about food or that other unrest is not also driven by food prices.

THEORETICAL FRAMEWORK

This paper poses the relatively straightforward research question: *Do sudden increases in local consumer food prices lead to an increase in the likelihood of social or political unrest of any type, violent or not, in African countries?* The innovation of this question is two-fold. First, the independent variable is increases in the prices that consumers actually pay for food on local markets rather than prices of internationally traded commodities. Second, the dependent variable is not limited to food riots but rather encompasses broad social unrest, including but not limited to labor strikes, student protests, electoral demonstrations and violence, and communal conflict.

It seems fairly obvious and well established that rising food prices *can* lead directly to food riots. There is much less agreement around whether rising food prices *will* lead to food riots and under what conditions. Do people simply respond to economic pressure or do they react to perceived injustices in the food production and distribution systems?

If people respond to economic pressure, we should to find a statistical correlation between consumer food prices and the occurrence of unrest. Therefore, it is not enough to look for an association between international commodity prices and unrest. Consumers do not pay international commodity prices and the economic pressure causal mechanism linking international commodity prices to unrest must necessarily function through the passthrough effect of international prices on local food prices.

Without identifying the link between commodity prices and consumer prices, any observed correlation between changes in the FAO Index and food riots, as observed by Lagi et al. (2011) and Bellemare (2012), putting aside the endogeneity problem for a moment, begs another explanation besides economic pressure from rising food prices. As Bush (2010) suggests, food riots could be the result of perceived injustices in the international food system regardless of any pass-through effect to local consumer food prices, but this is the result of a different causal relationship.

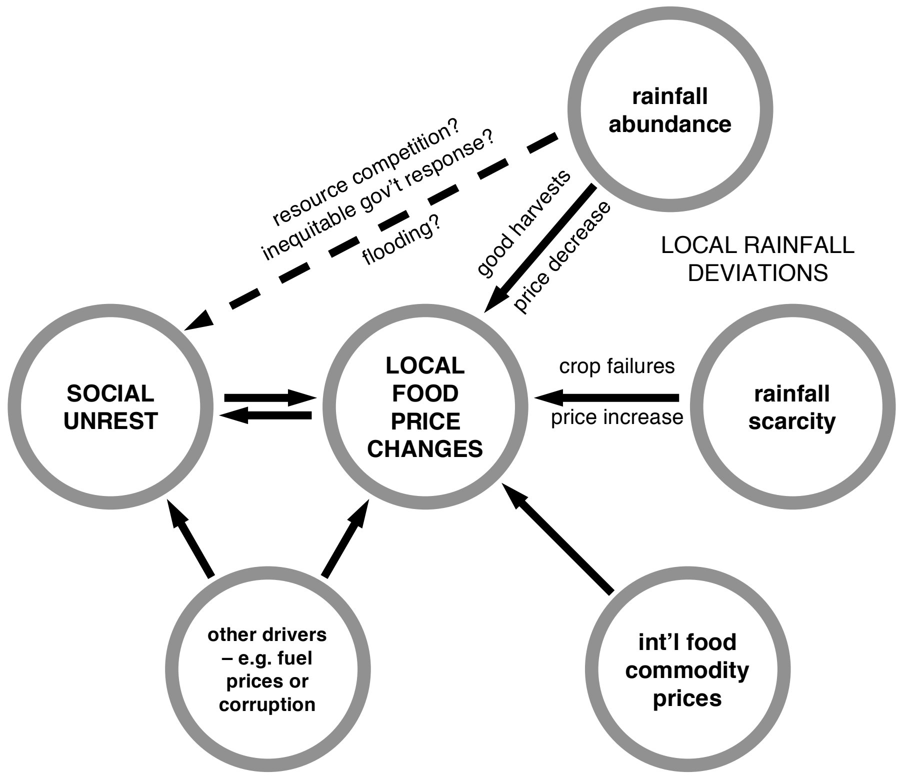
A number of potential problems prevent identifying a link between food prices and unrest. The first problem lies in the definition of the dependent variable, which has generally been confined to food riots. It is, however, unrealistic to expect populations to perceive their hardship in the same way, to frame their grievances in the same terms, or to address their petitions to the same audience. As discussed above the narrow category of food riots take on many different forms and are often enmeshed with other grievances, notably labor disputes.

In Africa, rising food prices have an undeniably disproportionate impact on poor populations who often devote over half of their income to food purchases (Kendall, Olson, and Frongillo 1996; Ruel, Haddad, and Garrett 1999; Maxwell et al. 2000) Economic hardships resulting from rising food prices could, however, manifest in many other forms. Even if people focus their protest on the government, they may demand economic relief in other sectors such as housing, utilities or fuel if they consider the government powerless to control rising food prices. Or they may seek a wholesale change in political leadership if they consider the current leaders to be incapable of creating favorable economic conditions. Furthermore, people may focus on a different target. As observed by Bush (2010) in Egypt, workers may strike with demands of higher wages that would enable them to afford more expensive food.

Although it is difficult to identify particular instances of unrest driven by rising food prices if the protesters do not cite food prices as a grievance, there are numerous examples of in which food is just one of many grievances. In Nigeria, the removal of fuel subsidies in December 2011 caused fuel and food prices to rise sharply, which led to widespread political protests that threatened the government of President Goodluck Jonathan. (Nossiter 2012a; Nossiter 2012b; Nossiter 2012c) In Uganda, protests ostensibly over rising food and fuel prices were intimately linked to perceived political oppression of opposition leaders. (Kron 2011) A protest in South Africa was organized by a leading labor union and cited rising electricity rates and food prices as grievances. (IRIN 2008)

Second, rising food prices and social unrest are both products of the same endogenous system. Not only do rising food prices cause social unrest, but social unrest, including food riots, can also drive up domestic food prices because of disruptions in the production or delivery of food supplies. Furthermore, it is possible that both rising food prices and social unrest are driven by other factors. This is what happened in Nigeria in December 2011 and January 2012. Rising fuel prices were likely the driver of both rising food prices and social unrest. This endogeneity can be the source of considerable bias in any attempt to empirically isolate the effect of rising food prices on social unrest, meaning that any statistical analysis that does not account for endogeneity will lead to incorrect conclusions. One method to eliminate this bias is to use an instrumental variable to predict the endogenous independent variable, in this case, food price spikes. Two potential instrumental variables are presented and evaluated: *international commodity prices* and *rainfall scarcity*. The theoretical framework is diagramed in Figure 2, which illustrates the bidirectional causal relationship of the dependent and independent variables, the potential effect of other variables, and how the instrumental variables function trough the independent variable or directly on the dependent variable.

Figure 2: Theoretical Framework



In order to function as a good instrument the chosen variable must must satisfy three assumptions: (1) it must not affected by either the independent or dependent variable; (2) it must be associated with changes in the independent variable; and (3) it must not lead to changes in the dependent variable, except indirectly through the independent variable. (Cameron and Trivedi 2010, 178)

The first instrumental variable I propose and test is international food commodity prices. These are the prices at which food commodities are traded on international markets and at international ports. Theoretically, the data on these prices meets the above criteria. First, it is reasonable to assume that international commodity prices are exogenous to local food prices in African countries. African countries are not significant producers of the commodities included in the

FAO index and, as such, are price-takers.[4] Second, national food prices are likely to be influenced by international commodities, especially in net food importing countries, which most African nations are to a lesser or greater extent. If the price of international food commodities rises, this is likely to be reflected in the prices of those goods in local markets. This assumption is easily testable. As has been previously discussed, the third assumption is more questionable. It is certainly true that any real economic pressure felt by local consumers as a result of international commodity prices will be through domestic food prices. The effect of any perceived injustices is more difficult to quantify. Another limitation of this as an instrument is that the monthly index price is constant for all countries despite the fact that it has different importance in different markets based on the type and amount of commodities imported.

The second, and ultimately better, instrumental variable is *rainfall scarcity*, that is lower than expected accumulated rainfall. This satisfies all of the assumptions of an instrumental variable. First, it is undeniably exogenous to the system. Neither food prices nor the occurrence of unrest can affect rainfall. Second, rainfall deviations have obvious impacts on food prices. Rainfall that does not come at the time or in the amount expected by farmers will lead to lower than expected harvests. Basic economics tells us that lower supply leads to higher prices. Conversely, better than expected rainfall may lead to bumper crops and a drop in prices. As discussed below, this is exactly what the data shows.

The third assumption is less obvious. Although, space does not permit a thorough discussion of the rainfall / conflict literature[5], it is increasing clear that there is some correlation between rainfall deviations and conflict. (Levy et al. 2005; Hendrix and Glaser 2007; Hendrix and Salehyan 2012; Raleigh and Kniveton 2012) The causal mechanisms remain unclear, however. In the absence of evidence of neo-Malthusian conflict over water resources, I argue that any connection between rainfall scarcity must be through agriculture, the most obvious effect of which is rising food prices. Rainfall abundance, however, is a very different phenomenon. While moderately good rains can have positive effects for agriculture, in the extreme, floods can have

[4] According to USDA estimates, only 7.25% of the world's total corn production in 2012 will be in African countries. The largest African producer is South Africa with an estimated 1.59% of the total in 2012 compared to 32.25% in the U.S. and 23.56% in China. (Index Mundi 2011)
[5] See Appendix B for a brief overview of the rainfall / conflict literature.

devastating effects. They can destroy infrastructure and displace populations, which can lead and has lead to conflict over scarce resources or unrest over inequitable government response. For these reasons, I argue that rainfall scarcity is an appropriate instrument but rainfall abundance is not because it affects conflict directly through different causal mechanisms.

ANALYSIS

This section will proceed as follows. I will first present my unit of analysis and the reasoning behind it. I will then present the details of the dependent and independent variables followed by a brief overview of the control variables. Next, I will present a discussion of the formulation, benefits, and limitations of the instrumental variables and a comparison of their validity as instruments in the tested model. This is followed by a presentation of the empirical analysis and the model results

The Unit of Analysis

I use a country-month unit of analysis, calendar months in individual African countries. This builds on the extant literature by using country level of geographic analysis of Arezki and Brückner (2011) combined with the monthly level temporal analysis of Bellemare (2012). This temporal disaggregation helps to better identify more immediate connections between food price changes and the occurrence of unrest events. Poor populations feel the pressure of rising food prices within days or weeks. If food prices are the proximate cause of unrest then the correlation should be observable within fairly short intervals of time.

Geographic aggregation to the national level allows for the use of fixed effects to control for country level characteristics that are likely to affect food prices and unrest. Characteristics such as natural endowments of arable land and weather conditions, internal transportation and food storage infrastructures, and access to external markets are all very important determinants of food prices. National trade and agricultural policies, such as tariffs and subsidies, have direct effects on both agricultural production and food prices. While country level aggregation sacrifices internal differences between rural and urban markets and political conditions, this is theoretically justified. While food prices may be driven by changes in growing conditions in rural areas, changes in food prices are more salient to urban populations who purchase a larger portion of their food. Protests are, therefore, more likely in urban centers. Furthermore, even when rural populations are aggrieved they may be inclined to take these grievances to the capital city where they will be heard.

The dependent variable: Social unrest

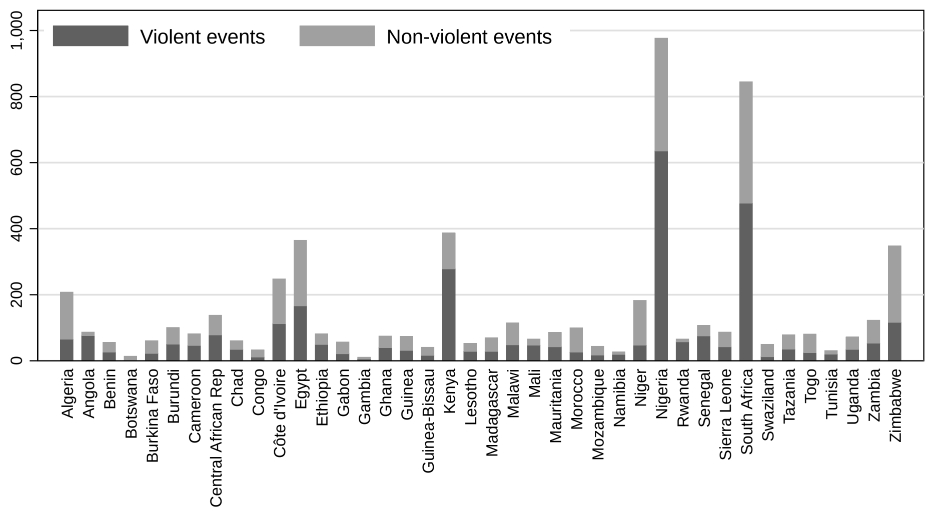
The dependent variable is the occurrence of social unrest, broadly defined, and is derived from the Social Conflict in Africa Database (SCAD) developed by Hendrix et al. (2010). SCAD includes data on over 7,300 occurrences of demonstrations, protests, strikes, riots, coups, and communal conflicts in 47 African countries[6] between 1990 and 2010. These events were compiled from *Associated Press* and *Agence France Presse* news reports, the major English and French language news services and are coded for event type – i.e. spontaneous or organized demonstration, riot, strike, pro- or anti-government violence, or extra- or intra-government violence – actors involved and targeted, and by motivating issue(s). Figure 3 shows the aggregate of these occurrences by country while the distribution of these events across time is presented later in Figure 5.[7]

Using the start date of each event I have recoded this data to create a binary variable for the occurrence of any unrest, violent or not, in a given country in a given month. Of the 11,804 country-months over the 21 years covered by SCAD, at least one unrest event occurred in 3,677 (30.72 percent) of them. In an effort to avoid normative judgments on the appropriateness or productivity of unrest, I created a second binary variable for the occurrence of *violent* unrest if any of the events in a given month were coded as violent. This is subset of the *unrest* variable

[6] African countries with a population of less than one million – Cape Verde, Comoros, Djibouti, Equatorial Guinea, São Tomé and Príncipe, and Seychelles – are excluded from SCAD.
[7] SCAD classifies events into ten different categories: (1) organized demonstrations; (2) spontaneous demonstrations; (3) organized violent riots; (4) spontaneous violent riots; (5) general strikes; (6) limited strikes; (7) pro-government violence (repression); (8) anti-government violence; (9) extra-government violence; (10) intra-government violence. Events in any category are considered to be unrest events. Violent events are limited to events in categories 3, 4, 7, 8, 9 and 10.

and occurs in 2,134 (18.18 percent) of the country-months covered by SCAD. The use of this dependent variable does not substantively change the conclusions and the results of this analysis are presented in Appendix E.

Figure 3: Social Unrest Events by Country (1990–2010)



Source: SCAD

Unfortunately, availability of other data limits the coverage of the SCAD data somewhat. As shown in Table 1, at least one unrest event occurred in 29.52 percent of the 9,365 country-months included in the final analysis. Violent unrest occurred in 16.52 percent of the observations.

The independent variable: National consumer food prices

The primary independent variable is national food price indices obtained from the International Labor Organization (ILO) for 47 countries between 1976 and 2012 (the panel is not complete across all years and countries). (International Labor Organization 2013) These data are reported to the ILO by each country and are collected for a basket of foodstuffs deemed to be most relevant in the capital city and other urban centers. Overlapping indices are reported for some years or some countries – Madagascar, Namibia. Sierra Leone, Swaziland and Zambia. In such cases, I choose the most complete index in order to avoid interruptions in the time-series. In none of these cases is the difference particularly pronounced except for Tanzania, in which case I opt for the mainland Tanganyika index and exclude the index for Zanzibar.

I then calculate the percentage change from one month to the next. This transformation has three distinct advantages. First, it better captures the immediate economic pressure felt by consumers of drastic short-term changes as opposed to more gradual changes to which consumers may be better able to adapt. Second, it enables valid comparison of food price changes across countries that use indices for prices of different products measured in different currencies. Third, it minimizes any statistical bias from autocorrelation over time that is inherent in price data of this type. In cases where the series was changed or reindexed, I exclude the percentage change for the first month of the new series in order to avoid any bias from an artificial change in food prices.

The instrumental variables

International commodity prices

I use FAO Food Index as the measure of international commodity prices. This index is “a measure of the monthly change in international prices of a basket of food commodities. It consists of the average of five commodity group price indices (representing 55 quotations), weighted with the average export shares of each of the groups for 2002-2004.” It is compiled and reported monthly with comparable data back to January 1990. The International Monetary Fund also compiles and reports an index of prices of food commodities. This index is highly correlated with the FAO index ($r = 0.9795$) but is available only beginning in 1991. Nonetheless, this alternative index is used as a robustness check. (For a comparison of FAO and IMF indices see Figure A3 in Appendix A.)

Rainfall scarcity

The rainfall indicator used here is based on data from the Global Precipitation and Climatology Centre (GPCC). (Schneider et al. 2013) This dataset includes monthly rain gauge data for the whole of the African continent since the beginning of twentieth century reported at a geographic resolution of 0.5 degrees latitude by 0.5 degrees longitude. Because of the lack of rain gauge stations in the first half of the century, the data prior to 1960 are questionable and I do not use them. The raw GPCC data was aggregated to monthly totals over each country in country in Africa. Ultimately, only the data between 1970 and 2010 are used.[8]

Similar to the calculation of the Standard Precipitation Index (SPI) developed by Guttman (1999), I then developed the new Moving Standardized Cumulative Precipitation (MSCP) over the preceding period, ranging from one to 18 months, for each country-month in the dataset. This I define as the cumulative rainfall of *n* months (current month plus the previous ***n-1*** months) minus the long-term moving average of the previous 20 years for the given calendar month divided by the long-term standard deviation of the previous 20 years.

The formula is written as follows:

$$MSCP_n = \frac{1}{\sigma}\left(\sum_{i=0}^{n} P_{t-(n-1)} - \mu\right)$$

P is monthly precipitation, μ is the long-term average over the preceding twenty years and σ is the long-term standard deviation of the same period.

Because the agricultural sector in each country is adapted to the climatic conditions within that country the same absolute changes are likely to have different effects in different countries. This transformation of standard deviations from expected has the benefit of inherently controlling people's expectation of rainfall and for rainfall conditions to which that the agriculture sector in each country is adapted. (Hendrix and Salehyan 2012; Koubi et al. 2012) Furthermore, using the observed patterns from previous twenty years, the memory of a single generation, as the long-

[8] An earlier version of this paper used data from the Global Precipitation and Climatology Project (GPCP) (Adler et al. 2003), which is a composite of three different sources, rain gauge data and two satellite sources, and has the benefit of greater global coverage and correction of potential errors than in any one data source. These data are only available from 1979 and are reported at a coarser geographic resolution of 2.5 degrees latitude by 2.5 degrees longitude. When aggregated to the country level these data are highly correlated (Pearson's correlation coefficient of 0.9672) with the GPCC data and did not significantly change the findings.

term average captures the expectations, particularly of farmers, for precipitation in the current year. This formulation addresses the concerns of Sovey and Green (2011) who criticize the use of rainfall as an instrumental variable for economic growth because local rainfall patterns may be predictable based on previous experience. Admittedly, it does little to address the potential cross-border effects that rainfall deviations in one country may have on the economic conditions of a neighboring country. I argue, however, that monthly changes in food prices are far less vulnerable to this violation of the independence assumption than annual economic growth. It is possible that the rainfall conditions may vary across some larger countries and the country aggregation will wash out these intra-country differences. Changes in food prices in the capital and other urban centers, however, are theoretically the sum of changes in local conditions across the country. Practically speaking, this aggregation is mandated by the unit of analysis and the other available data.

Finally, this single MSCP variable is bifurcated into two separate variables, dry MSCP and wet MSCP, which are the absolute values of negative and positive deviations respectively. This allows the two indicators to operate independently and can, therefore, capture the different mechanisms through which a scarcity or an abundance of rainfall may lead to changes in food prices and social unrest. In this way rainfall scarcity can function as an instrument independently of rainfall abundance, which may affect unrest through mechanisms other than food prices.

The effect of rainfall on food prices is inherently a product of timing. Normal rainfall in the two to three month growing season may compensate for rainfall scarcity in the remainder of the year. Additionally, there is likely to be a lag in between a shortfall in precipitation and a noticeable effect on the market. Nonetheless, I argue that an accumulated scarcity in rainfall over a period of several months, will likely have an evident effect on food prices. In order to select the appropriate period for accumulation of rainfall I separately estimated the fixed effect models of one through 18 months of MSCP and all covariates on the percentage change in national food prices, discussed in more depth later. I find that dry MSCP over a nine-month period has the most pronounced effect on food price changes with the narrowest confidence interval. (Figure B2 in Appendix B graphically presents the marginal effects and confidence intervals of three, six, nine, twelve, fifteen and eighteen months of dry MSCP.) The distribution of MSCP over nine months is roughly normally distributed (see Figure B3), a result of the imposition of normality by nature of the transformation. Summary statistics for both wet and dry MSCP are included in Table 1.

Table 1: Summary Statistics

	N	mean	sd	min	max
Occurrence of unrest	9,274	0.295	0.456	0	1
Occurrence of violent unrest	9,274	0.164	0.370	0	1
% change in national consumer food prices	9,274	0.939	3.266	-27.450	38.837
% change in FAO Food Commodity Index	9,274	0.333	2.488	-12.285	7.803
9 month dry MSCP	9,274	0.425	0.643	0	7.332
9 month wet MSCP	9,274	0.503	0.800	0	6.624
Occurrence of unrest (lagged)	9,274	0.295	0.456	0	1
National Elections	9,274	0.031	0.174	0	1
Population (millions)	9,274	17.836	24.078	0.855	160.092
Population Growth (monthly %)	9,274	0.198	0.106	-1.706	2.669
Urban Population (% of total)	9,274	35.803	16.557	5.392	86.034
Youth Population (% of total 14 & under)	9,274	42.637	4.667	23.338	49.146
GDP per capita (thousands of constant 2000 USD)	9,274	0.837	1.012	0.107	4.815
Polity IV democracy	9,274	2.786	3.081	0	9
Polity IV autocracy	9,274	2.746	2.654	0	10
Life expectancy at birth total (years)	9,274	53.310	8.001	26.819	74.600
Mortality rate infant (per 1000 live births)	9,274	78.709	28.304	14.8	162.0

Control variables

Consistent with Dunning (2011) and Linebarger and Salehyan (2012), I hypothesize that more unrest will occur in months in which national elections take place. I, therefore, include a single indicator variable for the occurrence of a presidential poll, legislative elections or a constitutional referendum in each country-month. This I developed using data from the African Elections Database (2012), and supplemented with separate research from individual country websites for each country in northern Africa.

I include variables for the relative institutional democracy or autocracy of countries because it is likely to affect the outbreak of unrest, although in ways that are more unpredictable. More democratic countries may be less vulnerable to social or political unrest if citizens believe they can make changes in their leadership and, thereby, in governmental policies through more formalized processes, such as elections or the court system. On other hand, people might feel more free to protest their government if their right to do so without governmental repression is protected by law. More autocratic countries will likely be less vulnerable to political or social unrest because the population may fear governmental repression. Although, if unrest does become manifest it is potentially more likely to be violently repressed by the government. In order to control for and measure these effects, I include variables from Polity IV Project for institutionalized democracy and autocracy separately, each of which is measured on a scale of zero to 10 with ten being the most democratic or autocratic respectively. (Marshall and Jaggers 2011)

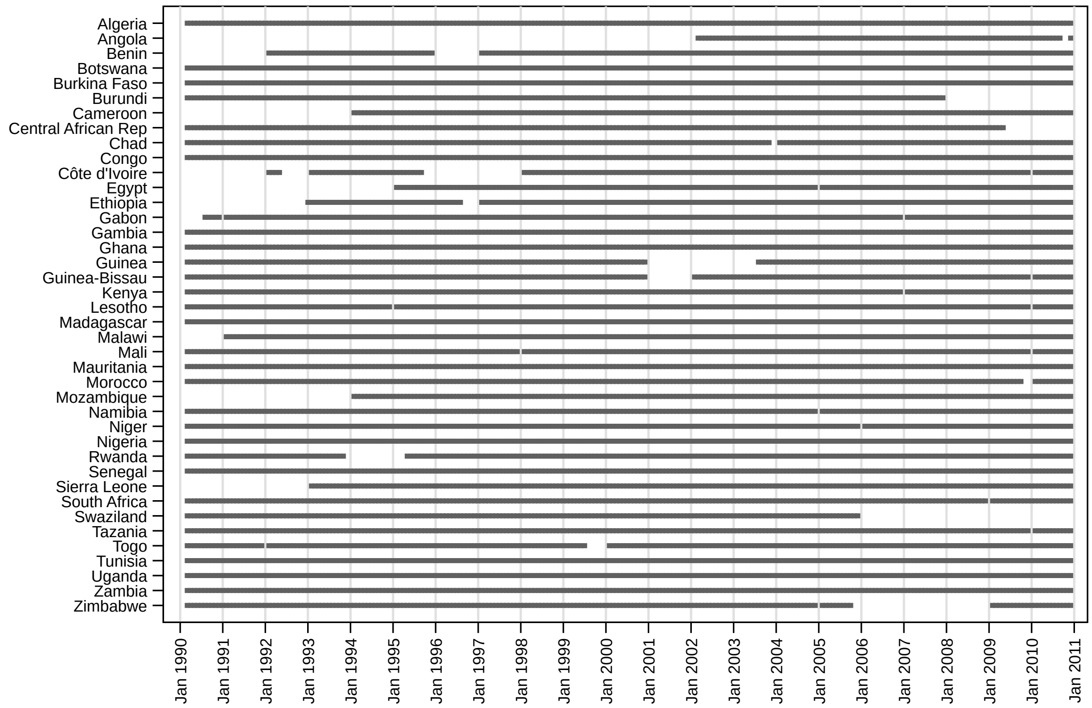
Demographic characteristics may also mitigate both how much rainfall deviations effect local food prices and how sensitive populations are to the rising food prices. I include variables for the size and growth rate of the population. Urban population as a percentage of total is included because urban dwellers are likely to purchase the vast majority of their food while rural populations may produce much of their own food. (Maxwell 1999; Frayne et al. 2009; Ziervogel and Frayne 2011) Size of youth population (ages 0-14) is included because a growing body of research has found an association between youth bulges and the likelihood of conflict. (Goldstone 2002; Urdal 2006)

The relative wealth or poverty of a population may also be associated with the likelihood of conflict in a given country. GDP or GDP per capita, however, does not capture inequality within a society. Countries that are wealthier but with higher levels of inequality are potentially more likely to experience conflict than poorer but more egalitarian societies. Unfortunately, good data on inequality is not available across time for most African countries. Additionally, official GDP figures do not adequately capture wealth in countries in which a large portion of economic activity occurs in the informal sector. For these reasons, I do not include GDP or GDP per capita in the primary models, although I do collect this data and include it in a robustness check. Rather, I include indicators of life expectancy at birth and infant mortality rates as proxies for the relative poverty or development of countries. All of these indicators were obtained from WDI. I also include country fixed effects to control for any other time-invariant unobserved heterogeneity between countries.

Finally, I include dummy variables for individual calendar months and individual years. The former is intended to control for any the seasonality of crop production and rainfall expectations the susceptibility of different societies to unrest during different times of the year. For example, Islamic societies may be less inclined to protest during Ramadan than at other times of the year. Year dummy variables should capture changes in global economic conditions and political changes.

The compilation of all these data resulted in a complete dataset of the primary variables of interest for 9, 274 country-months over 40 African countries between 1990 and 2010. This is an average of 231.85 country-months per country. Exact coverage of countries by month is shown in Figure 4. Table 1 presents the complete summary statistics of each of the variables used in the final analysis.

Figure 4: Sample Coverage by Country and Month



Estimation strategy

The model estimated is a *two-stage endogenous probit model*. The first stage is a model that predicts the percentage change in consumer food prices using the instrumental variable, dry MSCP, as well as all other covariates included the second stage. The first stage follows the general form:

$$FP_{it} = \beta_0 + \beta_1 dryMSCP_{it-1} + \beta_2 wetMSCP_{it-1} + \beta_3 FAO_{t-1} + \beta_4 SU_{it-1} + \sum \beta X_{it} + \sum \beta M + \sum \beta C + u$$

$MSCP_{it}$ (*wet* and *dry*) are the deviations from normal rainfall (as described above) in country *i* and month *t*; FAO_{t-1} is the percentage change in the FAO Food Index Price in month *t-1*; SU_{it-1} is a binary variable for the occurrence of social unrest in country *i* in month *t-1*; *X* is a vector of time varying control variables, including the occurrence of national elections and the Polity IV

Autocracy and Democracy scores; *M* is a vector of binary variables for each calendar month; *Y* is a vector of binary variables for each year; and *C* is a vector of country fixed effects.

The second stage is a ***probit*** model for that estimates the probability of the binary outcome, the occurrence of unrest. These models follow the general form:

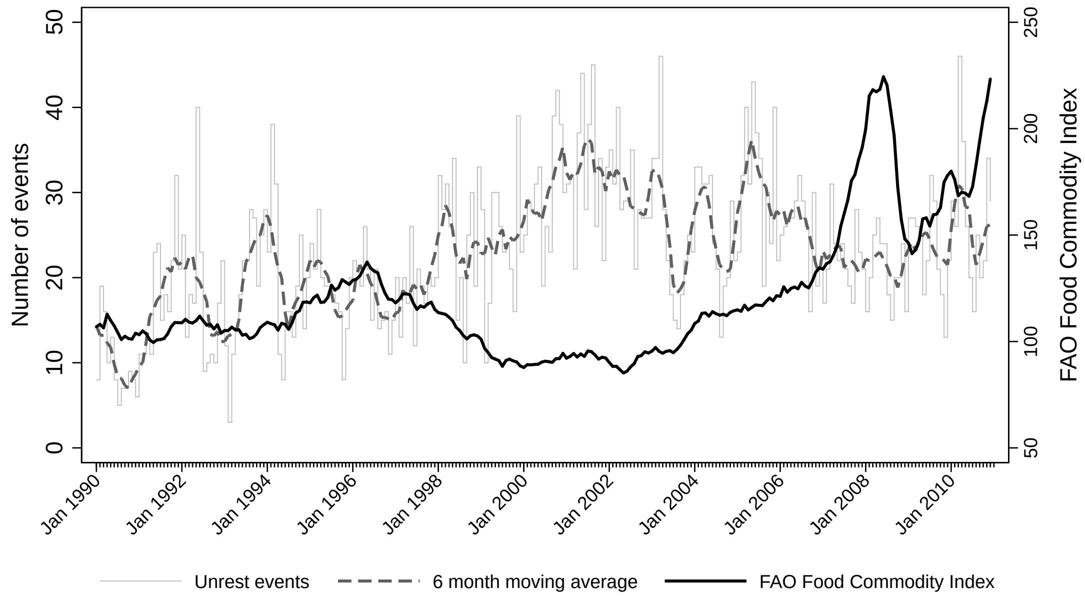
$$Pr(SU) = \Phi\left(\gamma_0 + \gamma_1 \widehat{FP}_{it} + \gamma_2 wetMSCP_{it-1} + \gamma_3 FAO_{t-1} + \gamma_4 SU_{it-1} + \sum \gamma X_{it} + \sum \gamma M + \sum \gamma C + (u + v)\right)$$

Pr(SU) is the probability of social unrest; Φ is defined as the normal cumulative distribution function; FP_{it-1} is the predicted percentage change in the food price index from the first stage equation and $MSCP_{it-1}$ (*wet* and *dry*), SU_{it-1}, *X*, *M*, and *C* are defined as above in the first-stage equation. As the instrumental variable, the *dry MSCP* is not included from the second-stage equation. The *ivprobit* routine is used to estimate the equation. (Newey 1987; Cameron and Trivedi 2010; StataCorp 2011) This is a maximum likelihood estimator that estimates coefficients for both equations in the same iterative process.

Results

An initial examination of the total number of SCAD events across the entire 21 years of data coverage, as shown in Figure 5, reveals no significant relationship between the two. In fact, it appears that the two move in opposite directions between 1998 and 2007. Drawing conclusions based on this is, however, not valid for reasons already discussed. More appropriate is a regression of the percentage change in national food prices and the occurrence of unrest in individual countries. The result of a fixed effects regression are presented in Table 2, models (1) and (2). This reveals that changes in food prices are marginally significant but the effects are very small and the fixed effects approach is not the best for the binary dependent variable. (Probit model results are presented in Table D2 of Appendix D.) The estimate of the effect of food price increases is biased by the endogeneity inherent in the relationship.

Figure 5: Social Unrest Events and International Food Commodity Prices



Univariate Prais-Winsten AR1 regression of FAO Food Price Index on number of SCAD events yields a coef. of -.012 (p = 0.6534)

The next step of the analytical process was to select the correct instrumental variable. In order to confirm that both international commodity prices and rainfall scarcity have the hypothesized impact on local consumer prices I specified a fixed effects model, the full results of which are presented in Appendix D. Although it is not the most appropriate model for a binary dependent variable, I used a fixed effects specification it allows for more thorough diagnostics of the instrumental variables that are not available in the other specifications. The relevant results of these models are presented in Table 2. I include the results for using change in the FAO index as the instrumental variable, model (2), dry MSCP at 9 months as the instrument, model (3), and both changes in FAO and dry MSCP as instruments, model (4). The Anderson canonical correlation test provides evidence that none of the models are underidentified. The Sargan-Hansen test for overidentification is only relevant for model (4) because the other two are exactly identified, but this test results in a failure to reject the null that the model is overspecified. {Baum, 2006}

The Cragg-Donald F statistic is a test for weak identification. This test statistic is compared against critical values established by Stock and Yogo (2002). Changes in the FAO index are shown to be a weak instrument. The test statistic of 6.59, is not greater than the twenty percent maximal IV size critical value meaning that considerable bias is likely still in the estimate of the instrumented variable, national consumer food price changes. (Stock and Yogo 2002)

Table 2: Fixed Effects Model Results (DV = occurrence of unrest)

	(1)	(2)	(3)	(4)	(5)
% change in national consumer food prices	0.002*	0.002*	0.019	0.082**	0.067**
	(0.001)	(0.001)	(0.050)	(0.035)	(0.028)
% change in FAO Food Commodity Index (lagged)	0.000	0.000	**INST VARIABLE**	-0.002	**INST VARIABLE**
	(0.002)	(0.002)			
9 month dry MSCP (lagged)	0.021***	0.021***	0.017	**INST VARIABLE**	**INST VARIABLE**
	(0.007)	(0.006)	(0.016)		
9 month wet MSCP (lagged)	0.005	0.008	0.010	0.016*	0.013*
	(0.007)	(0.006)	(0.008)	(0.009)	(0.008)
National Elections	0.187***	0.191***	0.190***	0.190***	0.191***
	(0.030)	(0.030)	(0.027)	(0.029)	(0.028)
Occurrence of unrest (lagged)	0.167***	0.155***	0.152***	0.137***	0.140***
	(0.016)	(0.016)	(0.017)	(0.016)	(0.015)
Unreported covariates	no	yes	yes	yes	yes
Country fixed effects	40 countries	40 countries	40 countries	40 countries	40 countries
Calendar month fixed effects	12 months	12 months	12 months	12 months	12 months
Year fixed effects	21 years	21 years	21 years	21 years	21 years
Observations	9274	9274	9274	9274	9274
Cluters	40	40			
Adjusted R-squared	.0552	.0645	.0427	-.314	-.182
Log psuedolikelihood	-4585	-4535	-4622	-6090	-5599
Anderson LR Statistic			6.62	20.4	27.4
Anderson LR chi2 p-value			.0101	.000006	.000001
Cragg-Donald F statistic			6.59	20.28	13.65
Stock-Yogo weak ID test crit values					
10% maximal IV size			16.38	16.38	19.93
20% maximal IV size			6.66	6.66	8.75
Hansen J-statistic			0	0	.976
Hansen J-statistic chi2 p-value					.323

Cluster robust standard errors reported for models (1) & (2); robust standard errors reported for models (3), (4) & (5); * p < .1; ** p < .05; *** p < .01; unreported covariates: Population; Population Growth; Urban Population; Youth Population; GDP per capita; Polity IV democracy; Polity IV autocracy; Life expectancy; Infant Mortality

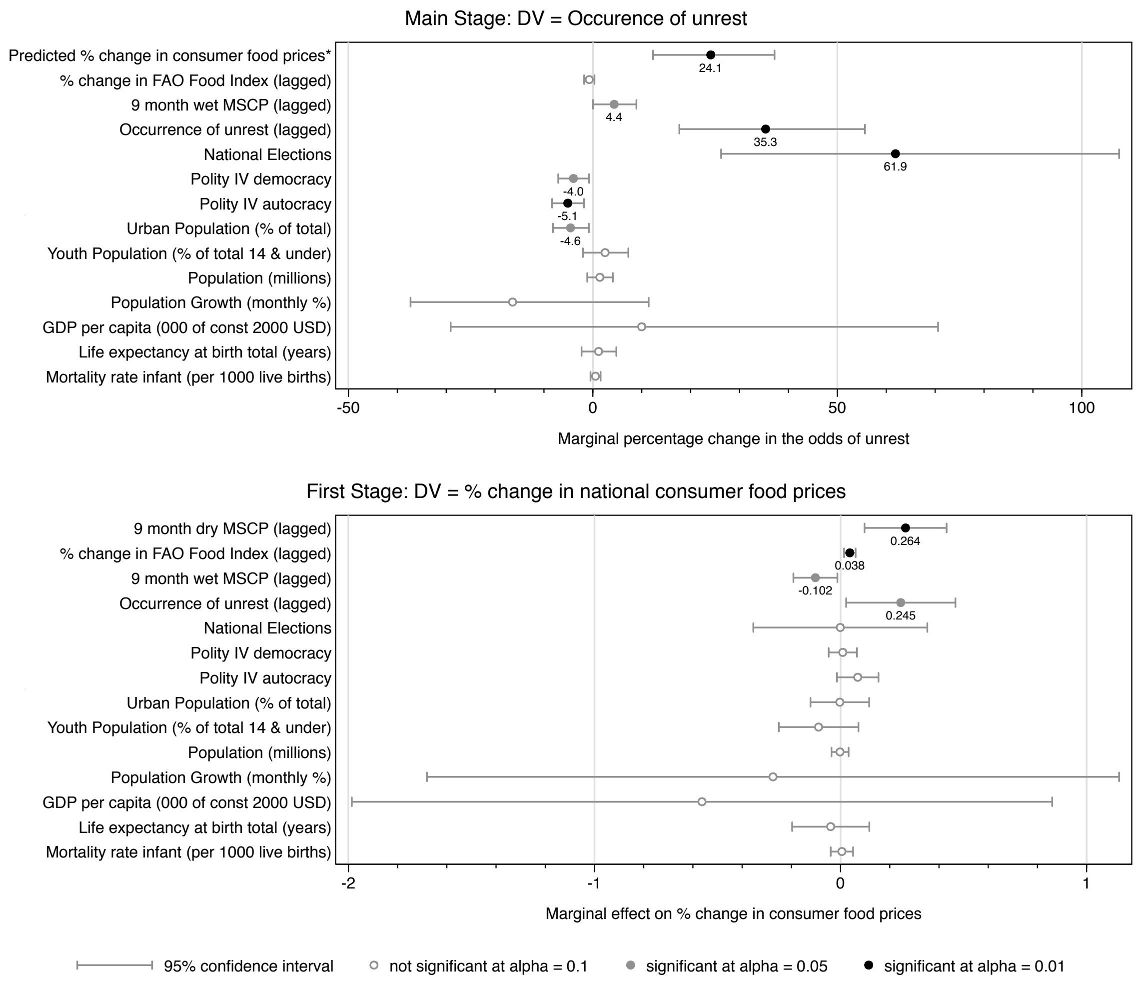
In sum, I conclude that changes in international commodity prices are not an appropriate instrument for changes in domestic food prices in their relationship with social unrest. This could be for a number of reasons. First, international food prices may lead to social unrest through another, as yet unidentified causal mechanism. This would explain the correlation observed by

Lagi et al. (2011) and Bellemare (2012), but it does not support their causal arguments. Second, because domestic markets are affected differently by changes in international market prices, it may be inappropriate to use the same data for every country in the sample. Countries import different commodities and in different quantities and have policies that control to different degrees how international changes are reflected in local prices. It may be possible in future research to account for these differences but that is beyond the scope of the current paper. Because FAO Index changes are a weak instrument on their own I also reject using both FAO changes and rainfall as joint instruments. The Cragg-Donald test statistic shows that the two together perform worse than rainfall scarcity on its own.

The results, however, do provide convincing evidence that rainfall scarcity, as measured by dry MSCP, is a good instrument for domestic food price changes. The test statistic of 20.28 is greater than the ten percent maximal bias critical value of 16.38. Furthermore, a Wald test, reported in Table D1 of Appendix D, supports the assumption of exogeneity. We can be satisfied that this is the best instrument for national food price changes in this model specification. This is the instrumental variable used in the next model presented.

The results of the main endogenous probit model previously described are presented graphically in Figure 6. (The full results of this model as well as models using the rejected instrumental variables are presented in Table D1 of Appendix D.) The first stage results are presented as marginal effects on the percentage change in consumer food prices. The direction of the effect of rainfall scarcity, change in FAO Index, rainfall abundance, and the occurrence of unrest in the previous month are all statistically significant, at least at the 95 percent confidence level, with effects in the hypothesized direction. are statistically significant at the 99% confidence level. All else equal, rainfall scarcity, accumulated rainfall one standard deviation below the long-term mean is associated with monthly consumer food price increases of 0.264 percent from the previous month (significant at alpha = 0.01). (Recall that both wet and dry MSCP are absolute values and, therefore, always positive.) Conversely, rainfall abundance, accumulated rainfall of one standard deviation above the long-term mean leads to a decrease in food prices of 0.102 percent. The effect of changes in the FAO Index on changes in national consumer prices is small but statistically significant (alpha = 0.01). On average, a one percent increase in the FAO index leads to an increase in national consumer food prices of 0.038 percent. Additionally, the occurrence of unrest in the previous month leads to an statistically significant (alpha = 0.05) increase in food prices of 0.245 percent. This is evidence of the endogenous nature of the relationship between food prices and unrest.

Figure 6: Endogenous Probit Model Results

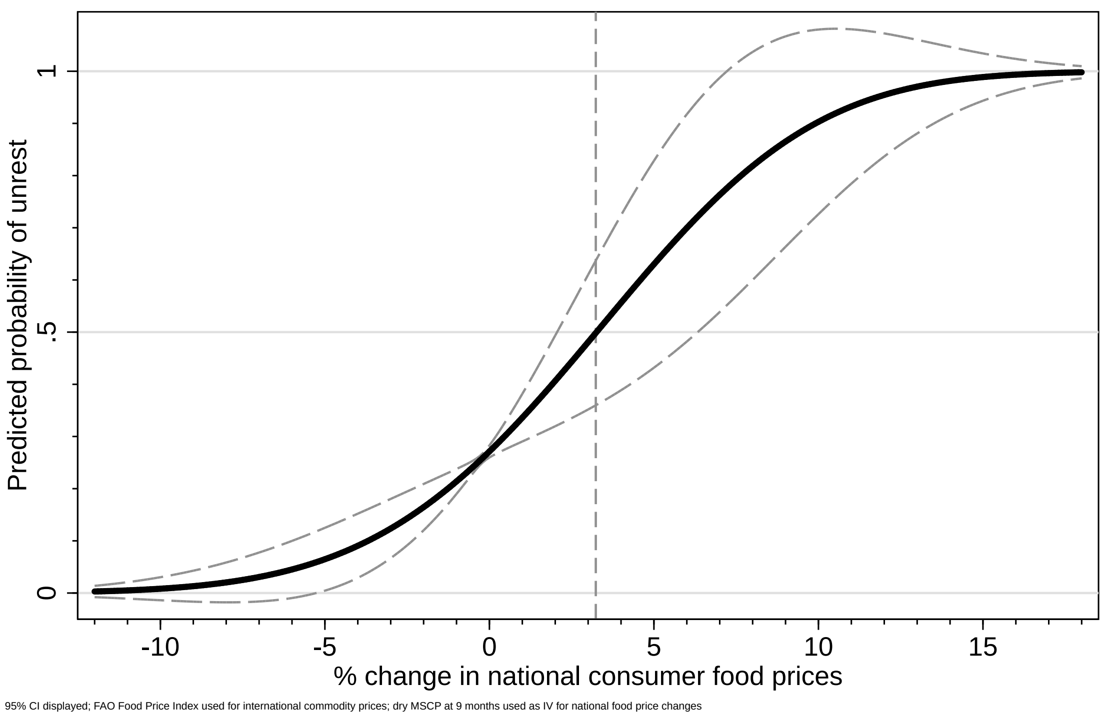


* % change in consumer food prices predicted using IV = 9 month dry MSCP (lagged) & all other covariates; country, calendar month & year fixed effects not displayed; 9266 observations; cluster robust standard errors employed (40 clusters)

The main stage results provide the answer to the central research question. Spikes in domestic consumer food prices do lead to increased probability of social unrest. The estimates are presented as marginal percentage changes in the odds of unrest. I estimate that each additional percentage point increase in the domestic consumer food price index will lead to a 24.1 percent increase in the odds of unrest. Not surprisingly, the occurrence of unrest in the previous month is a strong predictor of unrest. The odds of unrest in the current month are 35.3 percent higher if there was unrest in the previous month. National elections are also a strong predictor of unrest. The odds of unrest in an election month are 61.9 percent higher than a month without elections.

Rainfall abundance, although leading to lower consumer food prices, leads to a statistically significant (alpha = 0.05) increase in the odds of unrest. I interpret this as evidence that rainfall abundance effects unrest through a mechanism other than lower food prices. Some potential mechanisms were discussed briefly above but further examination of these possibilities is left for future research.

Figure 7: Effect of Food Price Changes on Predicted Probability of Unrest



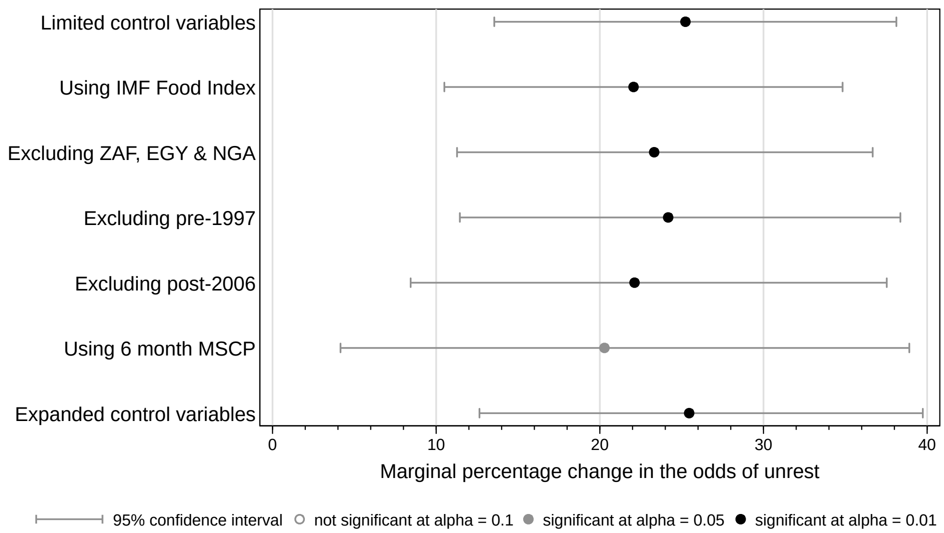
Although outside the scope of this paper it is also interesting that the odds of unrest are lower in countries with both higher levels of democracy and higher levels of autocracy. Countries in which a higher percentage of the population lives in urban areas are also less likely to experience unrest.

Understanding the magnitude of this effect is made difficult by the fact that changes in odds are dependent on the baseline odds. Figure 7 provides a better visualization of the effect of changes in the predicted probability of unrest. This graph shows that all else equal if food prices remain constant the probability of unrest is 27.1 percent. The probability of unrest is fifty percent, the odds are even, if food prices increase 3.23 percent in over the previous month. Food price increases greater than this mean that unrest is more likely than not.

Robustness checks

I performed a number of robustness checks, none of which significantly changed that findings for the change in consumer food price. First, I tested used a reduced set of covariates, only lagged unrest and national elections. Next, I used the IMF Food Index as the measure of international commodity prices in order to test for sensitivity to idiosyncrasies in the effect of the FAO index. Third, Egypt, South Africa and Nigeria may bias the results for a couple of reasons. Egypt and South Africa both have relatively developed irrigation infrastructure, which might make them less vulnerable to rainfall deviations. South Africa and Nigeria also account for a disproportionate amount of unrest events in SCAD, which might be a source of bias. I, therefore, exclude these countries from the sample to ensure that they were not biasing the results. Fourth, there is the potential of that more events are captured by SCAD after the use of the Internet became more widespread by media outlets. Although I consider the two sources used by SCAD, *Associated Press* and *Agence France Presse*, to be less prone to this potential source of bias than other less well-established outlets, I tested the robustness of the findings by limiting the sample to 1997 and later. Next, I tested whether the results are driven by the dramatic rise in international prices in 2007–08 and 2010, by excluding everything post-2006 from the analysis. Next, although I chose the appropriate rainfall measure with great care, I, nonetheless, tested the sensitivity of the estimates to different MSCP accumulations. MSCP at six months is presented. Lastly, I included additional control variables in an attempt to better capture the openness of the economy to international trade and the reliance of the national food system on internationally traded commodities.

Figure 8: Robustness Checks



As shown in Figure 9, the estimate for the effect of change in the consumer food prices is fairly stable across all models with odds ratios ranging from 1.203 in the MSCP 6 model to 1.255 in the model with the expanded set of control variables. Unsurprisingly, the change in rainfall accumulation period as the instrument has the greatest impact but the primary estimate of interest remains significant at the 95 percent confidence level. In all other models the estimate is significant at the 99 percent confidence level. (The full robustness model results of the robustness checks are presented in Appendix E.)

CONCLUSION

These results provide strong evidence of a causal relationship between changes in consumer food prices and the probability of unrest. Furthermore, they provide evidence that in the aggregate populations respond to the short-term economic pressure of rising food prices regardless of the cause of the increase. Basically, unrest is more likely at times of larger food price spikes. These results also demonstrate that changes in international commodity prices can lead and do increased probability of social unrest but the effect is, on average, minimal in Africa because the pass-through of commodity prices to local consumer prices is limited. Changes in local consumer prices are the real driver of general unrest.

Additionally, the finding that international commodity prices function poorly as an instrumental variable could indicate the presence of another causal mechanism linking international commodity markets to local unrest. This could be because of perceived injustices in the international system or it could be that both are the result of another factor. The price spike of 2007–08 and the observed food riots could both be, at least in part, consequences of the global economic crisis.

These findings have important policy implications. Price stability in local markets is important to maintaining political and social stability. Two important points to remember are when considering policy prescriptions are: (1) price spikes have destabilizing effects regardless of the source; and (2) consistently low prices are less important than preventing sudden spikes. Building stable local markets that are resilient to shocks from a variety of sources is key. Controlling price fluctuations in international markets is important at a macro level but will have little benefit if local markets are still vulnerable to local conditions. Uncertainty around climate change and changing weather patterns must be considered here but from a different perspective than it is usually considering in adaptation circles. Increasing overall long-term production through drought resistant crops or farming techniques is less important than ensuring availability and access in times of poor conditions. This means improving irrigation infrastructure to buffer against changing rainfall patterns, transportation networks to deliver food regionally, and storage capacity to allow for warehousing of food.

This paper cannot answer the theoretical question of how and why people choose particular methods, venues, and targets. More research is needed in this area but these questions are likely to be context specific. Answering these questions will require in depth case studies of particular episodes of unrest and are likely to have complex answers.

Just as the bread riots of the eighteenth and nineteenth centuries and the austerity protests of the 1970s and 1980s, episodes of unrest and political instability generally involve many issues are contributing factors. The list of possible drivers of unrest is long: perceived lack of justice and equity in global food and economic systems; lack of political freedoms and avenues to voice grievances; poor governance and lack of government accountability; general high levels of poverty and the lack of public service delivery; poor workplace conditions and labor disputes. All of these issues are intimately connected with access to food. To think that “food riots” are simply about food is a dangerous reduction. Such a view could easily lead one to focus on controlling the cost of food while ignoring political or social injustices. But it is equally blind to ignore that the cost of food can be a driver of unrest that manifests in other ways or is directed at targets seemingly unrelated to the food system.

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APPENDIX A: DOMESTIC AND INTERNATIONAL FOOD PRICES

Figure A1: National Consumer Food Indices for African Countries, Botswana – Lesotho

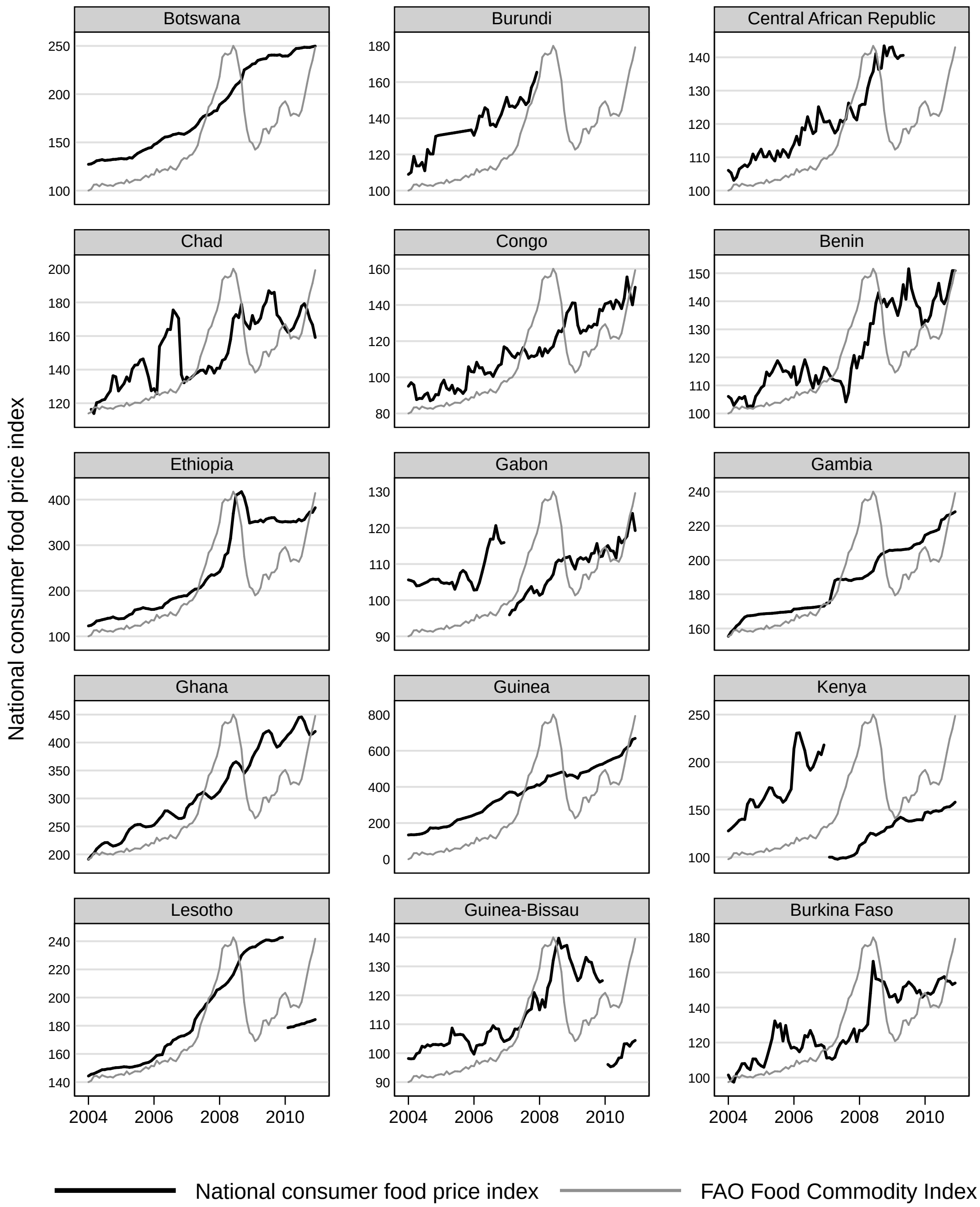
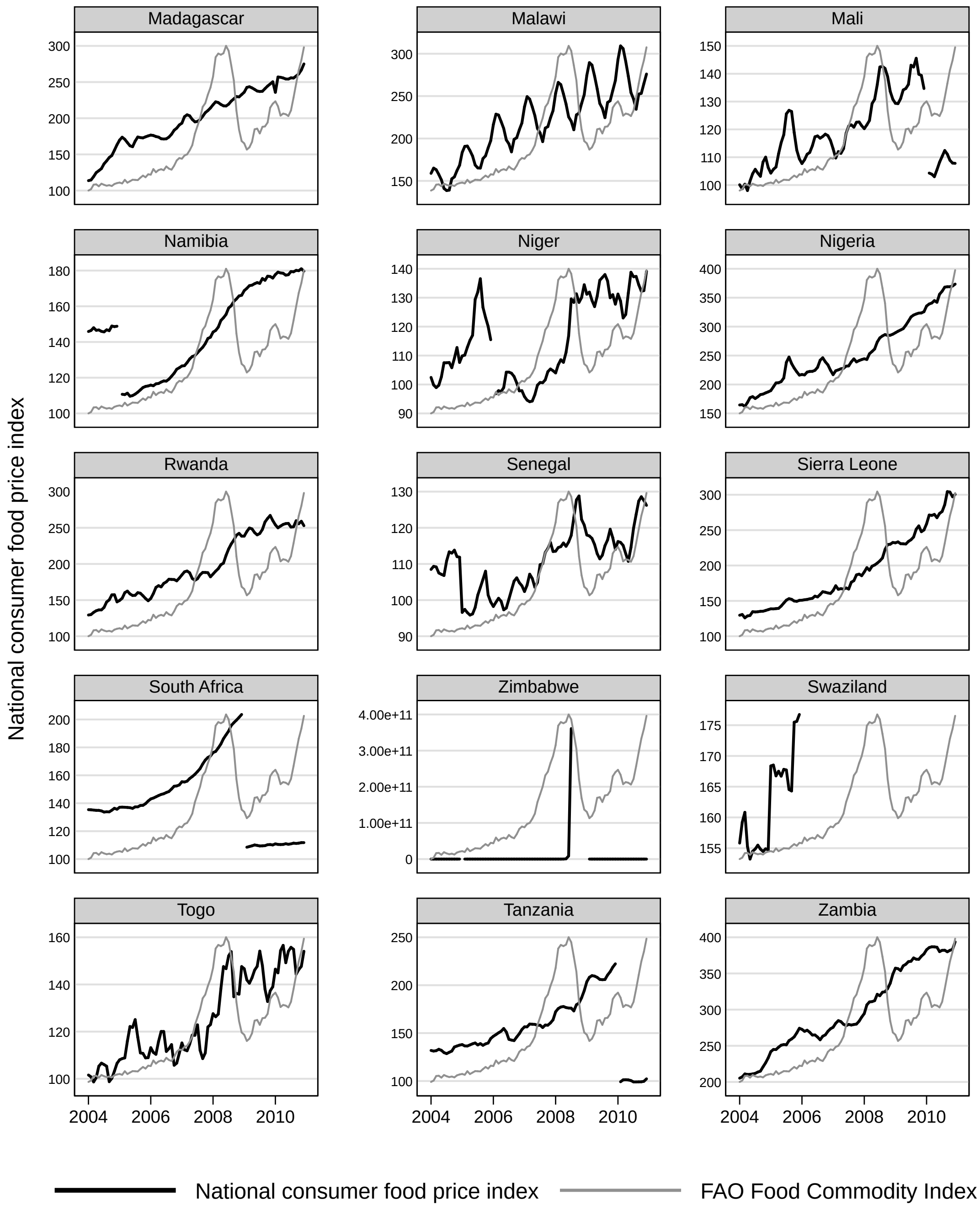


Figure A2: National Consumer Food Indices for African Countries, Mozambique – Zimbabwe



APPENDIX B: VARIABLE HISTOGRAMS

Figure B1: Histogram of Percentage Changes in Domestic Consumer Food Prices

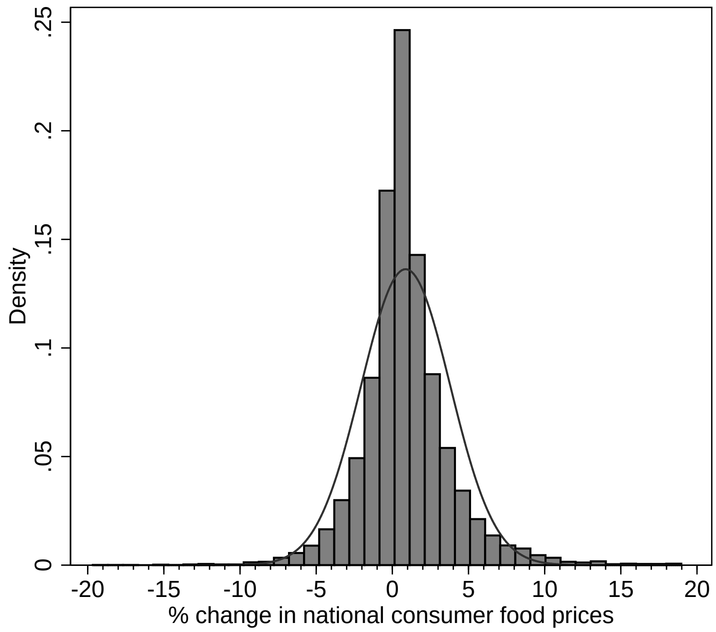


Figure B3: Histogram of Moving Standardized Cumulative Precipitation

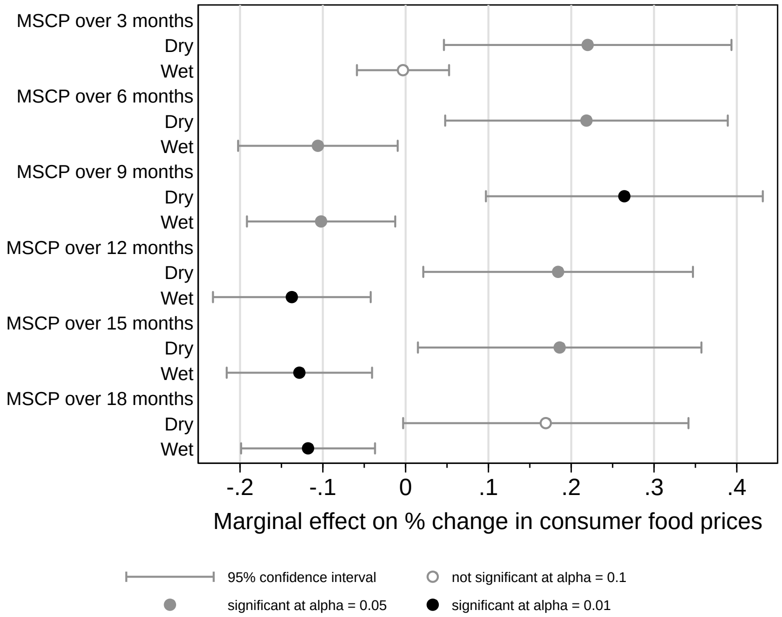
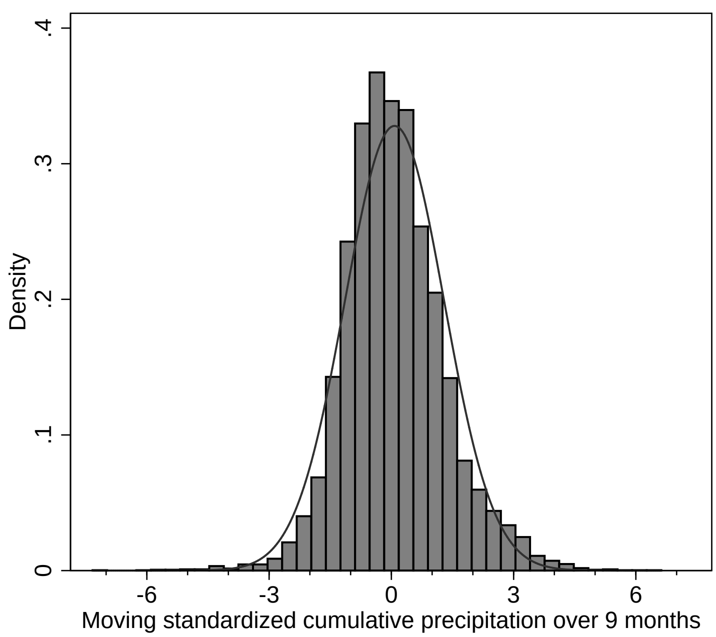


Figure B3: Histogram of Moving Standardized Cumulative Precipitation



APPENDIX C: RAINFALL AND CONFLICT LITERATURE

Several recent papers have examined the relationship between rainfall deviations and conflict using a variety of rainfall measures. Using a variety of definitions and sources of both conflict and rainfall extremes/deviations, some recent studies have found a relationship between rainfall deviations and conflict but the causal mechanisms remain largely unspecified. Attempts to shed light on the causal mechanism using large-N analysis have been largely inconclusive.

Using an annual grid square as the unit of analysis, Levy et al. (2005) conducted a logistic regression analysis of the effect of rainfall, measured with the Weighted Anomaly Standardized Precipitation index (WASP) aggregated annually, on the risk of civil war, using the Armed Conflict Dataset (ACD) from the International Peace Research Institute of Oslo and the University of Uppsala (PRIO/Uppsala). They find “a strong relationship between rainfall deviations, lagged one year, and the likelihood of high-intensity conflict outbreak” (Levy et al. 2005, p. 17). Hendrix and Glaser (2007) use the PRIO/Uppsala data and an annual time unit. They used the GPCP data to develop their independent variables: the percent change in rainfall from the previous year, which they called a rainfall trigger, and dummy variables for one and two standard deviations from mean annual rainfall. They find some evidence to support the hypothesis that “lagged measures of increased (decreased) rainfall are negatively (positively) associated with the risk of civil war onset”.

Using a grid square analysis, Theisen et al. (2011), however, find no evidence of a relationship between the occurrence of drought as defined using the Standardized Precipitation Index (SPI6), which “measures negative deviation from normal rainfall during the six preceding months” and the onset of armed conflict in the PRIO/Uppsala data (Theisen, Holtermann, and Buhaug 2011, p. 13). That droughts are not correlated with civil war, however, does not mean that drought cannot cause other lower level types of conflicts. Alternatively, the consequences of drought may be felt far from the actual occurrence of the drought or the measure of drought used in this analysis may not be relevant for some connections to conflict. They also performed robustness checks using a number of alternative drought measures.

Using SCAD, Hendrix and Salehyan (2012) find a relationship between annual standardized rainfall deviation, calculated as annual deviation from the long-term (1979–2008) mean divided by the long-term standard deviation and the number of social conflict events, defined more broadly than civil war or armed conflict to include strikes, riots and communal conflict. They further find that this relationship is strongest for violent events and positive deviations in rainfall have the largest effects (Hendrix and Salehyan 2012). In the same special issue, Raleigh and Kniveton (2012) find a significant linkage between monthly rainfall extremes, both wet and dry, and the occurrence of civil conflict in East Africa. They find that extreme wet periods are more associated with communal violence while extremely dry periods are associated with rebel conflict suggesting that different causal mechanisms are at work (Raleigh and Kniveton 2012).

Few studies have sought to empirically test different causal mechanisms. Building on a body of literature that suggests that economic growth is connected with civil war (Collier and Hoeffler 2002; Collier and Hoeffler 2004; Fearon and Laitin 2003), Miguel, Satyanath, and Sergenti (Miguel, Satyanath, and Sergenti 2004) use annually aggregated rainfall as measured by the Global Precipitation Climatology Project (GPCP) as an instrumental variable for economic growth. They find a strong negative correlation with the risk of civil conflict in Africa as measured by ACD. Ciccione (2011) has, however, criticized this analysis and found that Miguel et al.'s findings are "driven by a (counterintuitive) positive correlation between civil conflict and rainfall levels in year t – 2" (Ciccone 2011, p. 218). He concludes that more disaggregated data and analyses are necessary.

In a study similar to that of Miguel et al., Koubi et al. (2012) investigate the link between economic growth and civil war, defined as 25-battle related deaths in the Uppsala Conflict Data Program (UCDP) dataset, using rainfall deviations from normal as an instrument for economic growth. They find no association between rainfall deviations and economic growth, which indicates that the former is, at best, a weak instrument for economic growth in its relationship with conflict (Koubi et al. 2012). Bergholt and Lujula (2012) find limited evidence of a connection between economic growth and conflict using climatological natural disasters as an instrumental variable for economic growth, although the exogeneity of their natural disaster variable is highly questionable. They use the number of people affected by a storm or flood as reported in the EM-DAT database, but the number of people affected is determined not just by the occurrence of a storm or flood but also by how vulnerable the population is to such an event, which is highly dependent on economic development (Bergholt and Lujala 2012).

These seemingly disparate findings are not necessarily at odds. Grid square analysis suggests that rainfall and conflict are not spatially correlated at such a localized level. They do not, however, establish that a lack of rain in agricultural areas will not lead to conflict in urban centers. These large-N studies are consistent with many case studies that find limited or no evidence of a direct connection between drought and conflict amongst pastoralist groups (Adano et al. 2012; Butler and Gates 2012). The connection between rainfall, economic growth and civil war also seems to be questionable given the studies of Ciccione (2011), Koubi et al. (2012), and Bergholt and Lujula (2012). The observed connection between rainfall deviations and higher levels of conflict found by Hendrix and Salehyan (2012) and Raleigh and Kniveton (2012) still begs a causal explanation. Although the respective authors provide different hypotheses, I suggest that the causal connection might be through food prices.

APPENDIX D: FULL MODEL RESULTS

Table D1: Fixed Effects Model Results (DV = occurrence of unrest)

	(1)	(2)	(3)	(4)	(5)
% change in domestic consumer food prices	0.002*	0.002*	0.009	0.081**	0.063**
	(0.001)	(0.001)	(0.049)	(0.034)	(0.028)
% change in FAO Food Commodity Index (lagged)	0.000	0.000	**INST VARIABLE**	-0.003	**INST VARIABLE**
	(0.002)	(0.002)		(0.002)	
9 month dry MSCP (lagged)	0.021***	0.021***	0.019	**INST VARIABLE**	**INST VARIABLE**
	(0.007)	(0.006)	(0.015)		
9 month wet MSCP (lagged)	0.005	0.008	0.008	0.016*	0.012
	(0.007)	(0.006)	(0.007)	(0.009)	(0.008)
National Elections	0.187***	0.191***	0.191***	0.191***	0.191***
	(0.030)	(0.030)	(0.027)	(0.029)	(0.028)
Occurrence of unrest (lagged)	0.167***	0.155***	0.153***	0.136***	0.140***
	(0.016)	(0.016)	(0.017)	(0.016)	(0.015)
Population (millions)		0.004	0.004***	0.005***	0.005***
		(0.003)	(0.001)	(0.001)	(0.001)
Population Growth (monthly %)		-0.066	-0.064	-0.044	-0.051
		(0.054)	(0.058)	(0.065)	(0.061)
Urban Population (% of total)		-0.020***	-0.020***	-0.019***	-0.019***
		(0.005)	(0.003)	(0.003)	(0.003)
Youth Population (% of total 14 & under)		-0.002	-0.001	0.006	0.004
		(0.007)	(0.006)	(0.005)	(0.005)
GDP per capita (thousands of constant 2000 USD)		-0.030	-0.026	0.015	0.005
		(0.052)	(0.041)	(0.040)	(0.036)
Polity IV democracy		-0.014***	-0.014***	-0.015***	-0.015***
		(0.005)	(0.003)	(0.004)	(0.003)
Polity IV autocracy		-0.014***	-0.015***	-0.020***	-0.018***
		(0.005)	(0.005)	(0.005)	(0.004)
Life expectancy at birth total (years)		-0.001	-0.001	0.002	0.001
		(0.006)	(0.003)	(0.003)	(0.003)
Mortality rate infant (per 1000 live births)		0.002	0.002*	0.001	0.001
		(0.001)	(0.001)	(0.001)	(0.001)
Country fixed effects	40 countries	40 countries	40 countries	40 countries	40 countries
Calendar month fixed effects	12 months	12 months	12 months	12 months	12 months
Year fixed effects	21 years	21 years	21 years	21 years	21 years
Observations	9274	9274	9274	9274	9274
Cluters	40	40			
Adjusted R-squared	.0552	.0645	.0427	-.314	-.182
Log psuedolikelihood	-4585	-4535	-4622	-6090	-5599
Anderson LR Statistic			6.62	20.4	27.4
Anderson LR chi2 p-value			.0101	6.43e-06	1.13e-06
Cragg-Donald F statistic			6.59	20.28	13.65
Stock-Yogo weak ID test crit					
10% maximal IV			16.38	16.38	19.93
20% maximal IV			6.66	6.66	8.75
Hansen J-statistic			0	0	.976
Hansen J-statistic chi2 p-value					.323

Cluster robust standard errors reported for models (1) & (2); robust standard errors reported for models (3), (4) and (5); * p < .1; ** p < .05; *** p < .01

Table D1: Probit Model Results

DV = Occurrence of unrest	(6)	(7)	(8)	(9)	(10)
% change in domestic consumer food prices	1.008* (0.004)	1.008* (0.004)	1.028 (0.158)	1.241*** (0.063)	1.210*** (0.068)
% change in FAO food price index (lagged)	1.003 (0.005)	1.001 (0.006)	**INST VARIABLE**	0.993 (0.005)	**INST VARIABLE**
9 month dry MSCP (lagged)	1.078*** (0.026)	1.076*** (0.023)	1.070 (0.050)	**INST VARIABLE**	**INST VARIABLE**
9 month wet MSCP (lagged)	1.025 (0.025)	1.029 (0.024)	1.031 (0.031)	1.044** (0.023)	1.038 (0.025)
Occurrence of unrest (lagged)	1.667*** (0.084)	1.598*** (0.079)	1.589*** (0.114)	1.353*** (0.097)	1.404*** (0.100)
National Elections	1.849*** (0.168)	1.892*** (0.183)	1.890*** (0.188)	1.619*** (0.206)	1.686*** (0.214)
Population (millions)		1.018 (0.015)	1.018 (0.015)	1.014 (0.013)	1.016 (0.014)
Population Growth (monthly %)		0.731 (0.155)	0.735 (0.141)	0.836 (0.123)	0.809 (0.121)
Urban Population (% of total)		0.939*** (0.019)	0.939*** (0.019)	0.954** (0.019)	0.951*** (0.018)
Youth Population (% of total 14 & under)		1.008 (0.026)	1.010 (0.026)	1.025 (0.024)	1.024 (0.023)
GDP per capita (thousands of constant 2000 USD)		0.970 (0.239)	0.981 (0.247)	1.100 (0.246)	1.084 (0.246)
Polity IV democracy		0.950*** (0.016)	0.950*** (0.016)	0.960** (0.016)	0.958** (0.017)
Polity IV autocracy		0.951*** (0.018)	0.950** (0.021)	0.949*** (0.017)	0.948*** (0.017)
Life expectancy at birth total (years)		1.004 (0.022)	1.005 (0.024)	1.012 (0.018)	1.012 (0.018)
Mortality rate infant (per 1000 live births)		1.009* (0.005)	1.009* (0.005)	1.005 (0.005)	1.006 (0.005)
Country fixed effects	40 countries	40 countries	40 countries	40 countries	40 countries
Calendar month fixed effects	12 months	12 months	12 months	12 months	12 months
Year fixed effects	21 years	21 years	21 years	21 years	21 years
Observations	9274	9274	9274	9274	9274
Clusters	40	40	40	40	40
Log pseudolikelihood	-4482	-4427	-28139	-28139	-28139
Chi^2 model test			2148	5864	9058
Chi^2 model test p-value			0	0	0
Wald chi^2 test of exogeneity			.0168	9.8	6.74
Wald chi^2 test of exogeneity p-value			.897	.00174	.00942

Odds ratios reported; cluster robust standard errors reported in parentheses; * p < .1; ** p < .05; *** p < .01

Table D1: Probit Model First Stage Results

DV = % change in domestic consumer food prices	(8)	(9)	(10)
% change in FAO food price index (lagged)	0.038***	0.038***	0.031**
	(0.012)	(0.012)	(0.013)
9 month dry MSCP (lagged)	0.264***	0.264***	0.276***
	(0.085)	(0.085)	(0.082)
9 month wet MSCP (lagged)	-0.102**	-0.102**	-0.098**
	(0.046)	(0.046)	(0.046)
Occurrence of unrest (lagged)	0.245**	0.245**	0.244**
	(0.113)	(0.113)	(0.113)
National Elections	-0.001	-0.001	-0.001
	(0.180)	(0.180)	(0.180)
Population (millions)	-0.002	-0.002	-0.002
	(0.018)	(0.018)	(0.018)
Population Growth (monthly %)	-0.274	-0.274	-0.270
	(0.718)	(0.718)	(0.716)
Urban Population (% of total)	-0.003	-0.003	-0.003
	(0.061)	(0.061)	(0.061)
Youth Population (% of total 14 & under)	-0.089	-0.089	-0.090
	(0.083)	(0.083)	(0.083)
GDP per capita (thousands of constant 2000 USD)	-0.563	-0.563	-0.565
	(0.726)	(0.726)	(0.727)
Polity IV democracy	0.009	0.009	0.009
	(0.029)	(0.029)	(0.029)
Polity IV autocracy	0.070	0.070	0.070
	(0.043)	(0.043)	(0.043)
Life expectancy at birth total (years)	-0.040	-0.040	-0.040
	(0.080)	(0.080)	(0.080)
Mortality rate infant (per 1000 live births)	0.005	0.005	0.005
	(0.023)	(0.023)	(0.023)
Country fixed effects	40 countries	40 countries	40 countries
Calendar month fixed effects	12 months	12 months	12 months
Year fixed effects	21 years	21 years	21 years
Observations	9274	9274	9274
Clusters	40	40	40

Cluster robust standard errors reported in parentheses; * p < .1; ** p < .05; *** p < .01

APPENDIX D: USING VIOLENT UNREST DEPENDENT VARIABLE

Table E1: Fixed Effects Model Results (DV = occurrence of violent unrest)

	(1)	(2)	(3)	(4)	(5)
% change in domestic consumer food prices	0.002 (0.002)	0.002 (0.001)	-0.049 (0.043)	0.061** (0.028)	0.033 (0.022)
% change in FAO Food Commodity Index (lagged)	-0.002 (0.001)	-0.002 (0.001)	**INST VARIABLE**	-0.004** (0.002)	**INST VARIABLE**
9 month dry MSCP (lagged)	0.017*** (0.005)	0.016*** (0.004)	0.029** (0.013)	**INST VARIABLE**	**INST VARIABLE**
9 month wet MSCP (lagged)	0.003 (0.006)	0.005 (0.006)	-0.000 (0.007)	0.011 (0.007)	0.006 (0.006)
National Elections	0.171*** (0.030)	0.174*** (0.030)	0.174*** (0.026)	0.174*** (0.027)	0.174*** (0.026)
Occurrence of unrest (lagged)	0.093*** (0.014)	0.084*** (0.014)	0.097*** (0.015)	0.070*** (0.013)	0.077*** (0.012)
Population (millions)		0.006* (0.004)	0.006*** (0.001)	0.006*** (0.001)	0.006*** (0.001)
Population Growth (monthly %)		-0.059* (0.032)	-0.072* (0.044)	-0.042 (0.044)	-0.053 (0.038)
Urban Population (% of total)		-0.007* (0.004)	-0.008*** (0.002)	-0.007*** (0.002)	-0.007*** (0.002)
Youth Population (% of total 14 & under)		0.002 (0.004)	-0.002 (0.006)	0.007 (0.004)	0.005 (0.004)
GDP per capita (thousands of constant 2000 USD)		-0.068 (0.051)	-0.096*** (0.036)	-0.034 (0.031)	-0.049* (0.027)
Polity IV democracy		-0.011*** (0.003)	-0.011*** (0.003)	-0.012*** (0.003)	-0.011*** (0.003)
Polity IV autocracy		-0.013*** (0.004)	-0.009* (0.005)	-0.017*** (0.004)	-0.015*** (0.003)
Life expectancy at birth total (years)		-0.001 (0.005)	-0.003 (0.003)	0.002 (0.002)	0.001 (0.002)
Mortality rate infant (per 1000 live births)		0.001 (0.001)	0.001 (0.001)	0.001 (0.001)	0.001 (0.001)
Country fixed effects	40 countries	40 countries	40 countries	40 countries	40 countries
Calendar month fixed effects	12 months	12 months	12 months	12 months	12 months
Year fixed effects	21 years	21 years	21 years	21 years	21 years
Observations	9274	9274	9274	9274	9274
Cluters	40	40			
Adjusted R-squared	.0311	.0395	-.19	-.279	-.0531
Log psuedolikelihood	-2782	-2738	-3710	-4045	-3145
Anderson LR Statistic			6.85	20.4	27.7
Anderson LR chi^2 p-value			.00887	6.26e-06	9.85e-07
Cragg-Donald F statistic			6.82	20.3	13.8
Stock-Yogo weak ID test crit					
10% maximal IV			16.38	16.38	19.93
20% maximal IV			6.66	6.66	8.75
Hansen J-statistic			0	0	4.94
Hansen J-statistic chi2 p-value					.0262

Cluster robust standard errors reported for models (1) & (2); robust standard errors reported for models (3), (4) and (5); * p < .1; ** p < .05; *** p < .01

Table E2: Probit Model Results

DV = Occurrence of violent unrest	(6)	(7)	(8)	(9)	(10)
% change in domestic consumer food prices	1.006 (0.006)	1.007 (0.006)	0.822** (0.079)	1.247*** (0.056)	1.183** (0.080)
% change in FAO food price index (lagged)	0.992 (0.006)	0.990 (0.006)	**INST VARIABLE**	0.985*** (0.005)	**INST VARIABLE**
9 month dry MSCP (lagged)	1.089*** (0.024)	1.080*** (0.021)	1.119*** (0.028)	**INST VARIABLE**	**INST VARIABLE**
9 month wet MSCP (lagged)	1.017 (0.029)	1.024 (0.030)	0.998 (0.026)	1.040* (0.025)	1.029 (0.028)
Occurrence of unrest (lagged)	1.481*** (0.084)	1.418*** (0.078)	1.378*** (0.117)	1.229*** (0.073)	1.299*** (0.093)
National Elections	1.979*** (0.200)	2.039*** (0.211)	1.740*** (0.304)	1.696*** (0.220)	1.848*** (0.257)
Population (millions)		1.022* (0.012)	1.017* (0.009)	1.017* (0.010)	1.019* (0.011)
Population Growth (monthly %)		0.607** (0.146)	0.642 (0.178)	0.733 (0.147)	0.667* (0.158)
Urban Population (% of total)		0.956** (0.020)	0.965 (0.023)	0.968 (0.020)	0.962* (0.020)
Youth Population (% of total 14 & under)		1.013 (0.022)	0.992 (0.023)	1.029 (0.025)	1.027 (0.025)
GDP per capita (thousands of constant 2000 USD)		0.512* (0.177)	0.531* (0.194)	0.688 (0.188)	0.617 (0.197)
Polity IV democracy		0.948*** (0.013)	0.961** (0.018)	0.960*** (0.013)	0.954*** (0.014)
Polity IV autocracy		0.949*** (0.017)	0.974 (0.026)	0.948*** (0.016)	0.945*** (0.017)
Life expectancy at birth total (years)		1.011 (0.029)	1.001 (0.033)	1.017 (0.019)	1.018 (0.021)
Mortality rate infant (per 1000 live births)		1.009* (0.005)	1.008 (0.008)	1.006 (0.004)	1.007* (0.004)
Country fixed effects	40 countries	40 countries	40 countries	40 countries	40 countries
Calendar month fixed effects	12 months	12 months	12 months	12 months	12 months
Year fixed effects	21 years	21 years	21 years	21 years	21 years
Observations	9274	9274	9274	9274	9274
Clusters	40	40	40	40	40
Log pseudolikelihood	-4482	-4427	-28139	-28139	-28139
Chi^2 model test			2148	5864	9058
Chi^2 model test p-value			0	0	0
Wald chi^2 test of exogeneity			.0168	9.8	6.74
Wald chi^2 test of exogeneity p-value			.897	.00174	.00942

Odds ratios reported; cluster robust standard errors reported in parentheses; * p < .1; ** p < .05; *** p < .01

Table E3: Probit Model First Stage Results

DV = % change in domestic consumer food prices	(8)	(9)	(10)
% change in FAO food price index (lagged)	0.038***	0.038***	0.028*
	(0.012)	(0.012)	(0.015)
9 month dry MSCP (lagged)	0.264***	0.264***	0.281***
	(0.085)	(0.085)	(0.083)
9 month wet MSCP (lagged)	-0.102**	-0.102**	-0.096**
	(0.046)	(0.046)	(0.046)
Occurrence of unrest (lagged)	0.245**	0.245**	0.244**
	(0.113)	(0.113)	(0.113)
National Elections	-0.001	-0.001	-0.002
	(0.180)	(0.180)	(0.180)
Population (millions)	-0.002	-0.002	-0.002
	(0.018)	(0.018)	(0.018)
Population Growth (monthly %)	-0.274	-0.274	-0.268
	(0.718)	(0.718)	(0.715)
Urban Population (% of total)	-0.003	-0.003	-0.003
	(0.061)	(0.061)	(0.061)
Youth Population (% of total 14 & under)	-0.089	-0.089	-0.090
	(0.083)	(0.083)	(0.083)
GDP per capita (thousands of constant 2000 USD)	-0.563	-0.563	-0.565
	(0.726)	(0.726)	(0.727)
Polity IV democracy	0.009	0.009	0.009
	(0.029)	(0.029)	(0.029)
Polity IV autocracy	0.070	0.070	0.070
	(0.043)	(0.043)	(0.043)
Life expectancy at birth total (years)	-0.040	-0.040	-0.041
	(0.080)	(0.080)	(0.080)
Mortality rate infant (per 1000 live births)	0.005	0.005	0.005
	(0.023)	(0.023)	(0.023)
Country fixed effects	40 countries	40 countries	40 countries
Calendar month fixed effects	12 months	12 months	12 months
Year fixed effects	21 years	21 years	21 years
Observations	9274	9274	9274
Clusters	40	40	40

Cluster robust standard errors reported in parentheses; * p < .1; ** p < .05; *** p < .01