The Ethanol Mandate and Corn Price Volatility

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Abstract

Food price shocks are a major public concern, particularly in low income regions. This paper examines the impact of the United States ethanol mandate on food price volatility. We focus on corn price volatility because corn is both a major ethanol feedstock and a major international source of calories. Using agricultural commodity price time series, we provide suggestive evidence that a one billion gallon per year increase in the ethanol mandate increases corn price volatility by approximately 3.3 percent. Identification relies on a series of falsification tests which yield null results for commodities that are not directly related to ethanol production. Our results suggest that the ethanol mandate has increased the likelihood of very high price levels by even more than previously thought.

Keywords: Ethanol, biofuels, food price shocks, food security

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1 Introduction

Food price spikes in 2007-2008 and 2010-2011 caused mass hunger, exacerbated poverty in low-income nations, and prompted civil unrest (CNN (2008); Keating (2011)). Prior research has suggested that United States biofuel policy caused food prices to increase over this period (Zilberman et al. (2012); Wright (2014); Roberts and Schlenker (2010)). This paper asks whether U.S. biofuel policy has also made food prices more volatile. We use an econometric approach and provide evidence that one additional billion gallons of the ethanol mandate has increased corn price volatility by approximately 3.3 percent.

Congress passed the Energy Policy Act of 2005, creating the Renewable Fuels Standard (RFS). The RFS and subsequent RFS 2 mandate biofuel use in the U.S. These laws specify both a total mandated quantity and specific mandates for a variety of advanced biofuels. The total mandated quantity exceeds the sum of fuel-specific mandates. The difference can be met with a variety of fuels, but in practice corn-based ethanol is typically the cheapest and so the gap is commonly referred to as a corn ethanol mandate. The mandate has increased over time to its current level of approximately 14 billion gallons per year, which is more than ten percent of U.S. gasoline consumption.[1][2]

Approximately forty percent of the U.S. corn crop is now used to produce ethanol (USDA (2015)).[3] The mandate thus acts to shift corn demand outward dramatically. While estimates of the

[1]The EPA is to set particular targets each year in a fashion consistent with the RFS2, but due to delays the gap has not been well defined. To date, the EPA has limited RFS obligations to be less than 10% of total gasoline consumption. Furthermore, limited banking of compliance credits is allowed, so the actual compliance effort in any year may be somewhat less than the requirement if banked credits are available (and they have been). Nonetheless, ethanol consumption has increased dramatically under the mandate.

[2]Fuel is technically known as ”gasahol” when it contains ethanol. We use “gasoline” for familiarity.

[3]Some of this is returned to the food system as distillers’ grains and other coproducts.

precise magnitude vary, a number of studies have shown that this demand expansion has increased average corn prices (Fortenbery and Park (2008); Zilberman et al. (2012); Mitchell (2008); Roberts and Schlenker (2010); Carter et al. (2013); Wright (2014)).

Major agricultural commmodities are traded on exchanges with thick markets. Prices shift with shocks to supply, such as weather, or demand, such as income. Due to both complementaries and substitutability, many grain prices are cointegrated, and there are both price and volatility spillovers between commodities. A large literature explores this (Kim et al. (2009); Nazlioglu et al. (2013); Serra and Zilberman (2013)). However, few researchers focus on the direct relationship between the mandate and corn prices volatility. Econometric studies have shown that corn and energy prices are more tightly coupled since the mandate (Du et al. (2011); Serra and Zilberman (2013)). Simulation models find that the ethanol mandate increases corn price volatility, but provide some evidence that this is primarily due to an increased effect of agricultural supply shocks (Hertel and Beckman (2011); McPhail and Babcock (2012); Roberts and Tran (2013)). Wright (2011) attributes recent price shocks to a variety of supply, demand, and inventory factors of which the biofuel mandate was one. To our knowledge, this paper is the first to complement these theoretical and simulation studies with econometric evidence of the net effect of the mandate on corn price volatility.

Multiple models predict a change in corn price volatility due the ethanol mandate. One, detailed in Appendix A, is that the corn supply curve is convex. An increase in demand moves suppliers to a steeper part of the supply curve, exacerbating any shocks. A different model is a competitive storage model in which a mandate which increases over time causes storage operators to hold grain in anticipation of higher future prices. This means that less grain is available to buffer market shocks, increasing volatility. A calibrated model in Roberts and Tran (2013) shows that this can be a substantial effect.

The rest of this paper consists of following sections. Section 2 describes our econometric approach, while section 3 describes our data. Results are presented in section 4, while section 5 concludes.

2 Econometric models & tests

We will test for an increase in corn price volatility under the ethanol mandate. The core identification challenge is that there is a single trending treatment, so any treatment effect will not be well identified in the presence of nonlinear time effects. Therefore we instead adopt a falsification test approach which runs the same analysis for a wide variety of agricultural product prices. If the mandate does make corn prices more volatile, we would expect to see a treatment effect for corn. We would also expect to see a positive treatment effect for commodities that are closely linked with corn via substitutability in supply or demand - in particular sugar, which is closely linked via ethanol markets. We might also expect a treatment effect on wheat. Corn and wheat are substitutes in both production and use, and prior literature has found that price shocks to corn spillover to wheat (Gardebroek et al. (2014)). We would not expect to see an effect on other agricultural commodities. We estimate equation (1) for cotton, oats, rice, soybean, soybean meal, sugar and wheat

$$y_t = \mu + \beta \cdot mandate_t + X_t\gamma + \theta \cdot trend + \phi_m + \eta_{dow} + \varepsilon_t \tag{1}$$

where y_t is the outcome variable, $mandate_t$ measures the ethanol mandate, X_t is a vector of control variables, $trend$ is a linear time trend, ϕ_m is a month-of-year effect, η_{dow} is a day-of-week fixed effect , and ε_t is an unobserved error term. The error term ε_t is allowed to be autocorrelated of degree 1 such that $\varepsilon_t = \rho\varepsilon_{t-1} + \nu_t$ where ν_t is iid. Price volatility y_t is measured as the absolute value of change in logged prices $|\Delta \log(price_t)|$. This is the magnitude of daily price changes in percentage terms, so a larger value will indicate higher price volatility. The same transformation is applied to control variables X_t. We additionally use heteroskedasticity-robust standard errors.

The key testable hypothesis from our theoretical model is that a change in the mandate will change the level of corn price volatility.[4] If this is the case, then $\beta \neq 0$. However, if commodity markets broadly exhibit a nonlinear trend in volatility, then the estimated $\hat{\beta}$ may reflect time trends. Therefore we will not just examine the estimates. We will also test that the estimated $\hat{\beta}$ for corn is greater than and significantly different than the estimated $\hat{\beta}$ for other commodities. If it is, that will provide evidence that corn price volatility increased under the mandate relative to volatility in other agricultural commodities.

Both Phillips-Perron and Augmented Dickey-Fuller unit root tests show that y_t and X_t are stationary for all commodities. For all commodities, p-values are less than 0.0001. This is consistent with commodity prices which follow a random walk because y_t and X_t describe price volatility and not price levels and thus would not be expected to exhibit unit roots. This is robust to allowing

[4] Except the border case in which $D_P - S_P = D'_P - S'_P$, in which case the mandate will not change price volatility. This occurs if we assume a linear model.

an endogenously timed structural break.

3 Data

We use agricultural commodity prices from financial exchanges to calculate y_t. This data covers trading days from 1987 through 2013. Among independent variables, crop prices for corn, soybean and wheat are Chicago Board of Trade (CBOT) futures settle prices (front month) and in US dollars per bushel. Sugar prices are Intercontinental Exchange (ICE) No. 11, considered a world market price. We also use three economic variables to control market conditions: (1) Dow-Jones industrial average indices (DJ), a proxy for economic conditions; (2) West Texas Intermediate (WTI) spot prices (USD per barrel), which enables us to take into account that energy prices are related to the input cost of crop production[5]; and (3) a weighted average of the foreign exchange value of the US dollar against a subset of the broad index currencies (ER).[6] All price-related variables such as grain prices and WTI are depreciated using the monthly US Consumer Price Index to 2010 dollars (or cents).

Grain prices are set by trading in on financial markets in a rational expectations framework (Gorton and Rouwenhorst (2006)). Participants in grain markets include suppliers, consumers, market makers and speculators, and storage operators. Grain storage operators can purchase at low prices and sell at high. This permits consumers to access agricultural commodities outside the harvest season and links prices intertemporally (Deaton and Laroque (1992)). As long as

[5]Prices of crude oil and its products have bearing on prices of fertilizer, diesel-fuled farming equipments, transportation costs, etc.

[6]Major currencies index includes the Euro area, Canada, Japan, United Kingdom, Switzerland, Australia, and Sweden.

inventories are sufficient, storage operators can buffer positive price shocks by releasing inventories from storage. Storage operators can also buffer negative shocks by purchasing bulk grains. Within this supply, demand, and storage framework, grain price dynamics are consistent with trading in a rational expectations framework.

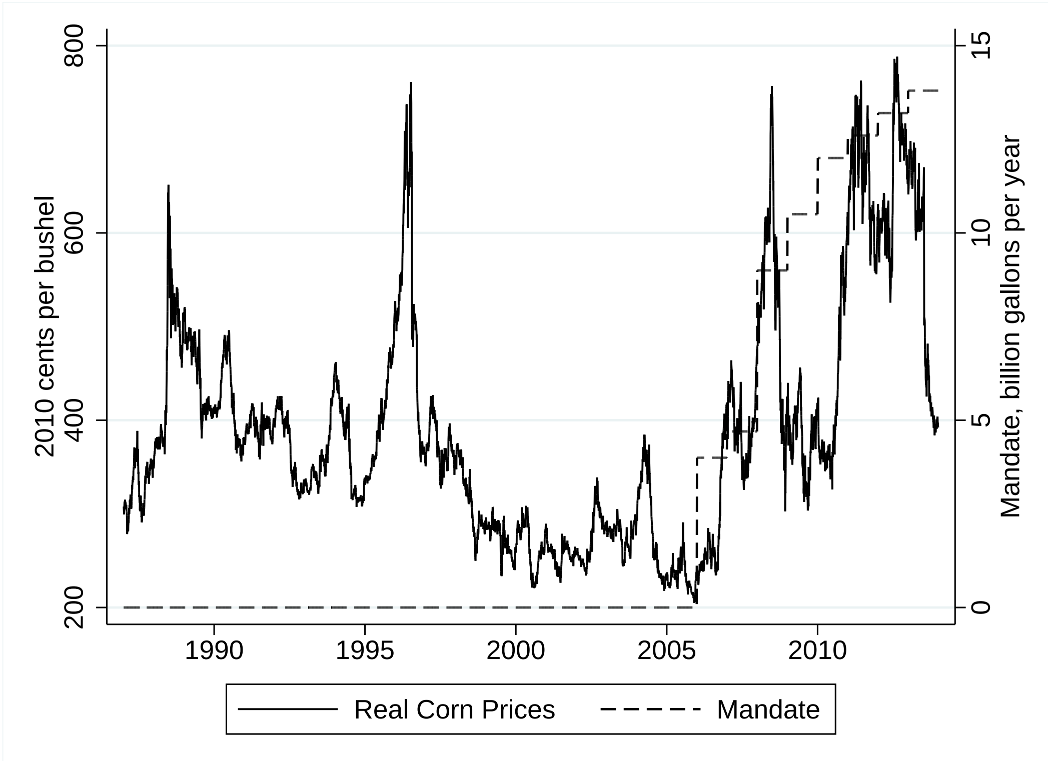
Ethanol mandate levels are measured in billions of gallons per year.[7] Obligated parties comply with the mandate via ethanol credits called Renewable Identification Numbers (RINs) which are both tradable and bankable. Forward looking agents will hold bankable allowances to smooth allowance prices over time.[8] Actual quantities of ethanol production may also be smoothed. To control for this smoothing, we will use both the actual annual mandate volume (shown in Figure 1) and a linearly smoothed mandate measure. The linearly smoothed measure is equal to the actual annual value on July 1 and is linearly interpolated on other dates.

Table 1 presents both price levels and y_t before and during the mandate. We see from the columns of mean absolute value of first difference in logs that the average daily percent change in corn prices before the mandate is 1.043%, but after it is 1.604%. We also see that other commodities have a mixture of increases and decreases in price volatility, but no other crops have a volatility increase as large as corn. We additionally see that corn prices increased under the mandate from an average of 349.5 cents per bushel before the mandate to an average of 484.4 cents per bushel under it and that the standard deviation of corn prices has increased under the mandate from 89.89 to 147.5. This is consistent with previous literature finding that the ethanol mandate caused an increase in corn prices and the hypothesis that it made prices more volatile, although dispositive

[7] The reader can find mandate levels in Table 1 of Schnepf and Yacobucci (2013). Actual effective mandated levels have differed due to EPA rulemaking. As the EPA rules were not always available during compliance periods, we use the statutory levels.

[8] For a full discussion of the structure of the RFS and economics of RINs, see Lade et al. (2015).

Figure 1: Corn Prices and the Mandate



of neither. Note that the number of observations is smaller by one for the first difference in the pre-treatment period - this is of course because we do not observe the change from day prior for the very first day of our study period.

Table 1: Descriptive statistics

	Before the Mandate						During the Mandate					
	Levels			Absolute value of first difference in logs			Levels			Absolute value of first difference in logs		
VARIABLES	N	mean	sd	N	mean	sd	N	mean	sd	N	mean	sd
Corn	4,791	349.5	89.82	4,790	1.043	1.102	2,009	484.4	147.5	2,009	1.604	1.529
Cotton	4,791	93.13	28.03	4,790	1.206	1.289	2,009	77.45	29.89	2,009	1.483	1.418
Oats	4,791	221.1	74.39	4,790	1.504	1.676	2,009	295.1	64.11	2,009	1.600	1.590
Rice	4,791	10.83	3.606	4,790	1.172	1.336	2,009	13.53	2.594	2,009	1.172	1.062
Soybean meal	4,791	273.8	72.77	4,790	1.146	1.199	2,009	319.9	78.37	2,009	1.440	1.450
Soybean	4,791	883.4	228.1	4,790	1.035	1.080	2,009	1,110	267.1	2,009	1.266	1.376
Sugar	4,791	13.79	4.695	4,790	1.571	1.630	2,009	17.75	5.805	2,009	1.693	1.638
Wheat	4,791	488.0	132.5	4,790	1.193	1.228	2,009	634.0	155.9	2,009	1.737	1.528

These data are presented visually in Figure 1, which shows real corn prices and mandate volume over time. We see that the mandate has increase from 0 before 2006 to a maximum of nearly 14 billion gallons per year. We also see that corn prices were broadly stable for the period immediately before the mandate, but have steadily increased under it. Visually, we also see large swings up and down in corn prices from 2006 onwards. Now we turn to testing these appearances more formally.

4 Results

Our results support the hypothesis that the ethanol mandate increased corn price volatility. We see positive estimates of the mandate on corn price volatility across specifications. These results are substantial, significant, and of consistent magnitude. Moreover, these results are statistically greater than estimates for all other crops except sugar. As sugar prices are closely linked to corn prices, this corroborates the core result.

The results of estimating equation (1) are presented in Tables 2–4. First consider Table 2 in which we estimate equation 1 with no control variables (i.e., no X variables). In column 1, we see that a one unit (one billion gallon) increase in the mandate level is associated with an increase of 0.0331 in the average daily percent price change. Referring back to Table 1, the average daily percent price change before the mandate was 1.0 (an average daily change of one percent per day), so a one billion gallon increase in the mandate level is associated with a 3.3 percent[9] increase in corn price volatility.

[9] Calculated as 0.0331 / 1.0.

Table 2: Results of estimation: testing for increase in volatility. Control variables omitted

	Corn	Cotton	Oats	Rice	Soybean	Soybean Meal	Sugar	Wheat
Mandate	0.0331***	0.0037	0.0106	-0.0145*	-0.0008	0.0118	0.0413***	0.0123*
	(0.006)	(0.006)	(0.008)	(0.006)	(0.006)	(0.006)	(0.007)	(0.006)
Constant	0.6896***	0.3794*	1.4962***	0.5562***	0.5357***	0.7465***	2.5912***	0.2681*
	(0.122)	(0.148)	(0.185)	(0.167)	(0.141)	(0.145)	(0.213)	(0.131)
Observations	6,799	6,799	6,799	6,799	6,799	6,799	6,799	6,799
R-squared	0.069	0.021	0.030	0.013	0.036	0.036	0.015	0.033

*Notes: *,**,*** indicate statistical significance at the 5%, 1%, and 0.1% level, respectively. Standard errors in parentheses. Results for month of year and day of week fixed effects are omitted.*

We additionally see positive point estimates for sugar and wheat which are significant at the 5% level, although only the effect on sugar is similar in magnitude or p-value to that on corn. Based on the same calculation as above, this also implies an increase in price volatility of approximately 2.5 percent for sugar and 1.0 percent for wheat and soybean meal. Sugar and corn compete in a variety of markets. In particular, Brazilian ethanol and U.S. ethanol compete in international markets. As ethanol and refined sugar can both be produced from the sugarcane in Brazil, we would expect a strong price relationship between corn and sugar. Furthermore, the prices are linked domestically. Sugar and corn are substitutes as sweeteners and previous researchers have found that their prices are closely related.[10] Thus the increase in demand for corn may have also increased demand for sugar and thus sugar price volatility.[11] Similarly, wheat substitutes with corn in a variety of domestic markets, and prior literature has found evidence of volatility linkages between the two (Gardebroek et al. (2014)).

Adding control variables in Table 3, point estimates and statistical significance remain largely unchanged.

[10]ICE argue that the sugar trade linked sugar prices to prices of corn-derived ethanol, and hence to the price of corn itself. Rapsomanikis and Hallam (2006) also report that sugar and ethanol prices are linearly cointegrated. In the context of the sweetener market, Moss and Schmits (2002) contend that there exists a substitution between high fructose corn syrup and sugar although the cointegration is not observed after 1996.

[11]Landes (2010) note that other concurrent changes in sugar markets may have independently increased sugar price volatility. In particular, they argue that the full implementation of NAFTA in 2008 newly exposed domestic sugar producers to volatile world markets, while a variety of idiosyncratic shocks caused prices to spike in 2009-2010.

Table 3: Results of estimation: testing for increase in volatility

	Corn	Cotton	Oats	Rice	Soybean	Soybean Meal	Sugar	Wheat
Mandate	0.0331***	0.0031	0.0121	-0.0153*	-0.0005	0.0110	0.0397***	0.0154**
	(0.006)	(0.006)	(0.008)	(0.006)	(0.006)	(0.006)	(0.007)	(0.006)
Dow-Jones	6.8368***	6.3632***	7.5871**	6.9497***	5.0971*	4.3045*	6.3026*	9.9424***
	(1.939)	(1.908)	(2.645)	(1.769)	(2.317)	(1.963)	(2.775)	(1.983)
Exchange rate	21.2454***	25.5896***	24.5954**	21.6808***	27.8180***	18.7777***	21.8977**	21.1451**
	(5.975)	(5.755)	(7.571)	(5.642)	(6.629)	(5.519)	(7.067)	(6.441)
WTI	4.9440***	2.4641**	4.5177***	2.9184*	4.2691***	2.3548**	2.6374*	5.3334***
	(1.007)	(0.870)	(1.321)	(1.239)	(0.994)	(0.912)	(1.180)	(1.227)
trend	0.0000**	0.0001***	-0.0000	0.0001***	0.0001***	0.0001***	-0.0001***	0.0001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Constant	0.4777***	0.1992	1.3195***	0.3775*	0.3538*	0.6100***	2.4070***	0.0800
	(0.124)	(0.146)	(0.188)	(0.170)	(0.139)	(0.145)	(0.215)	(0.130)
Observations	6,754	6,754	6,754	6,754	6,754	6,754	6,754	6,754
R-squared	0.081	0.030	0.040	0.022	0.050	0.041	0.019	0.050

*Notes: *,**,*** indicate statistical significance at the 5%, 1%, and 0.1% level, respectively. Standard errors in parentheses. Results for month of year and day of week fixed effects are omitted.*

Table 4: Results of estimation: testing for increase in volatility with smoothed mandate

	Corn	Cotton	Oats	Rice	Soybean	Soybean Meal	Sugar	Wheat
Mandate	0.0334***	0.0028	0.0122	-0.0167**	-0.0014	0.0109	0.0382***	0.0142*
	(0.006)	(0.006)	(0.008)	(0.006)	(0.006)	(0.006)	(0.007)	(0.006)
Dow-Jones	6.8374***	6.3654***	7.5879**	6.9599***	5.1024*	4.3059*	6.3125*	9.9518***
	(1.937)	(1.908)	(2.644)	(1.769)	(2.318)	(1.963)	(2.773)	(1.983)
Exchange rate	21.2212***	25.6008***	24.5900**	21.7420***	27.8507***	18.7779***	21.9435**	21.1928**
	(5.970)	(5.754)	(7.571)	(5.643)	(6.628)	(5.518)	(7.064)	(6.443)
WTI	4.9409***	2.4612**	4.5161***	2.9093*	4.2629***	2.3522**	2.6193*	5.3201***
	(1.007)	(0.871)	(1.321)	(1.238)	(0.994)	(0.912)	(1.180)	(1.227)
trend	0.0000**	0.0001***	-0.0000	0.0001***	0.0001***	0.0001***	-0.0001***	0.0001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Constant	0.4902***	0.1950	1.3230***	0.3481*	0.3386*	0.6110***	2.3886***	0.0618
	(0.125)	(0.147)	(0.189)	(0.173)	(0.140)	(0.146)	(0.218)	(0.132)
Observations	6,754	6,754	6,754	6,754	6,754	6,754	6,754	6,754
R-squared	0.081	0.030	0.040	0.022	0.050	0.041	0.019	0.050

*Notes: *,**,*** indicate statistical significance at the 5%, 1%, and 0.1% level, respectively. Standard errors in parentheses. Results for month of year and day of week fixed effects are omitted.*

No other crops have point estimates that are statistically greater than zero at conventional confidence levels, which is important because of our falsification test approach. The null result for other crops suggests that unobserved and broad commodity market effects were not the cause of the positive estimates for corn. Moreover, we can reject the null hypothesis that each commodity's β equals corn's β for all commodities except sugar. Table 5 presents results of t-tests of the hypothesis that each commodity's β equals corn's β. While we cannot reject the null that they are equal for sugar (p=.237), we can for all other commodities. In other words, while the estimated effect of the mandate on crop price volatility is not exactly zero, the mandate's effect on corn volatility is statistically larger than on other agricultural goods. This bolsters the prediction that the mandate increased price volatility of ethanol-related commodities. The t-tests presented are based on the results of Table 3, which are consistent with results from other models.

Table 5: Hypothesis testing on coefficients of mandates

	Coefficient of mandate	p-value (for the t-test)
Corn	0.0331	-
Cotton	0.0031	0.000***
Oats	0.0121	0.017**
Rice	-0.0153	0.000***
Soybean	-0.0005	0.000***
Soybean meal	0.0110	0.004***
Sugar	0.0397	0.237
Wheat	0.0154	0.018**

Notes:,**,*** indicate statistical significance at the 10%, 5%, and 1% level, respectively.*

As a robustness check, Table 4 presents results for estimating equation 1 with a linearly smoothed measure of the mandate. Point estimates and statistical significance are largely unchanged from either model using the annual mandate level.

Note that R^2 values are low across all estimates - typically less than 0.05. This is of course consistent with the efficient markets hypothesis that changes in commodity prices are difficult to predict. The explanatory power is modestly higher for corn prices, although still less than 0.1. This again supports the hypothesis that the mandate explains the magnitude of price changes for corn more than for other crops.

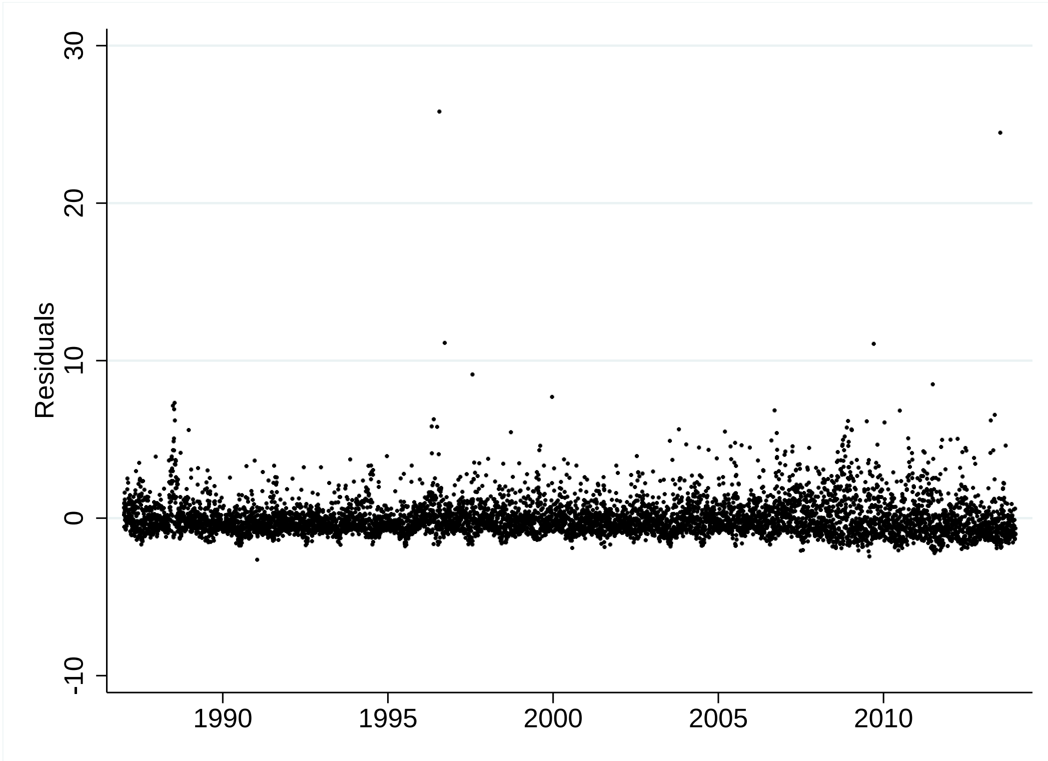
Finally, we consider the validity of our linear trend formation. Figure 2 shows the residuals from estimating equation 1 with controls and annual mandate level but without the time trend. There is no visual evidence that would contradict the use of a linear trend before our treatment period.

5 Conclusion

The Renewable Fuels Standard and RFS 2 were designed to promote energy security and reduce greenhouse gases from gasoline consumption. However, as with all policies, benefits should be balanced against costs - both potential environmental costs and spillovers into other markets. Agricultural price volatility can be a substantial burden on households, particularly households in low-income communities. A number of authors have attributed recent food shock (at least in part) to the U.S. ethanol mandate (Wright (2014)).

This paper provides suggestive evidence that the mandate has increased price volatility, which

Figure 2: Residuals over time



could lead to a higher likelihood of high realizations of prices. Further research continuing to explore the determinants of price volatility would hold substantial value.

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A Conceptual Model

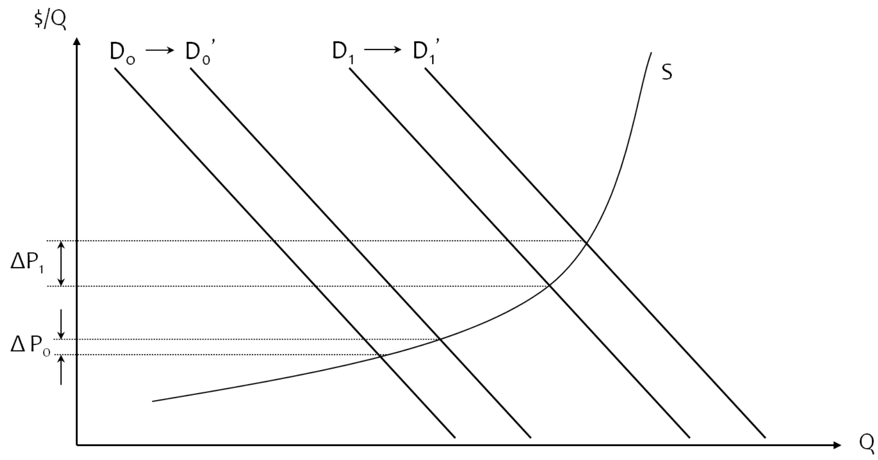
In this Appendix we develop a simple model of supply and demand for corn. The ethanol mandate acts to shift the demand for corn outwards.[12] Both supply and demand curves are vulnerable to exogenous shocks in their shifters. We show that these shocks could lead to higher or lower changes in prices for high mandate levels depending on the relative curvatures of the supply and demand functions. The intuition is that the supply curve may get steeper, but the demand curve will get shallower (if the mandate demand is not perfectly inelastic). A steeper supply curve will lead to a larger price change in response to a shock, but a shallower demand curve will lead to a smaller one. The net effect is thus ambiguous in sign and an empirical question. The model is depicted graphically in figure 3 and developed analytically below.

Corn demand comes from two sources. Baseline prior demand describes non-RFS corn demand in which corn is largely used for livestock feed, sweeteners, exports, and other food products, as well as as a minor energy source. The ethanol mandate adds a large amount of demand for transportation fuels both because of direct compliance obligations (which as noted above may vary from the statutory mandate level) and because it has prompted the dramatic expansion of ethanol refining capacity. Transportation fuel demand is quite price inelastic (Hughes et al. (2008)). The total demand is then the horizontal sum of these two demand sources, or alternatively one can consider the total demand to be shifted out by the ethanol mandate. This is shown in Figure 3 by the shift from D to D'.

Our analysis will be at daily frequency, so corn planting and harvest decisions are predeter-

[12] For analytic simplicity, we omit consideration of storage. Storage would act to buffer shocks, particularly negative price shocks, but prices in storage models still depend on the supply curve.

Figure 3: Conceptual Model



mined and inelastic from the perspective of our model. However, due to stock-holding behavior and corn's substitutability in other markets, the residual supply of corn to ethanol refiners is upwards sloping. It is not directly affected by the mandate, but can be affected an exogenous shifter.

Consider a simple supply-demand system in which both supply and demand depend on price and exogenous shifters α and β, respectively.

$$S = S(P, \alpha) \tag{2}$$

$$D = D(P, \beta) \tag{3}$$

We also assume market clearing: $S = D = Q$, where Q is the quantity of ethanol supplied and sold. We can rewrite this as

$$Q - S(P, \alpha) = 0 \tag{4}$$

$$Q - D(P, \beta) = 0 \tag{5}$$

By applying the implicit function theorem, we see that effect of a supply shock (ie, a change in α) on equilibrium price is

$$P_\alpha = \frac{S_\alpha}{D_P - S_P} \tag{6}$$

Now consider adding the ethanol mandate to the model via horizontal addition of the demand curves. This increases demand

$$D'(P, \beta) > D(P, \beta) \tag{7}$$

(where ' denotes the presence of the mandate) and flattens the demand curve

$$D'_P(P, \beta) \geq D_P(P, \beta) \tag{8}$$

The inequality is strict if mandate demand is not perfectly inelastic. It also moves up the supply curve, which has an ambiguous effect on S_P. If $S(P, \alpha)$ is convex, then $S'_P(P, \alpha) > S_P(P, \alpha)$.

Generally, the mandate will raise the response of prices to shocks if $D_P - S_P > D'_P - S'_P$. If $S(P, \alpha)$ is convex, then the sign is ambiguous - the mandate will increase volatility if increasing steepness of the supply curve outweighs flattening of the demand curve (and vice versa).