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Abstract

On May 29, 2013, the U.S. government announced that contamination of non-genetically-engineered (GE) wheat supplies occurred. We provide a methodology to empirically assess the impacts of this contamination event on national prices and on farm-specific prices and receipts for spring red hard wheat. Results suggest that U.S. spring red hard wheat farmers in the aggregate had lost receipts ranging from \$32.77 million to \$131.06 million and incurred a drop of 3.83% in wheat price equivalent to \$0.27 per bushel. At the farm level, a *hypothetical* farmer received \$0.31 less per bushel culminating in \$4,807 in lost receipts.

Key Words:

Assessment of national and farmer specific damages, May 2013 wheat contamination, use of vector error correction model genetically-engineered wheat

JEL Classifications: C39, C51, Q11

A Methodology to Empirically Assess National and Farm-Specific Damages from Contamination of Grain Supply by a Genetically-Engineered Strain

Introduction

Generally, much of the foreign demand for U.S. crops is constituted by notable importing nations who are adverse to accepting genetically-modified or GM crops and food products made with such crops because of perceived health risks to their environment and human health of their populations (Carter and Smith, 2007, p. 522; Li et al., 2010, p. 31). Quite simply, when the United States has, on a number of occasions noted below, announced that crop supplies have been contaminated with GM crop varieties or strains that have not been approved for human consumption, world demand for U.S. crops has fallen significantly and led to declines in U.S. farm prices, farm receipts, and farm income. There have been a number of studies that have estimated the U.S. national or “sector-level” crop price effects of such contamination events (see for example Carter and Smith, 2007; Li et al., 2010). Yet to date, we have been unable to locate studies that provide a methodology to take such national or sector-level effects and translate them into adverse decreases in farm-specific prices received, cash receipts and income for use by plaintiffs in litigation that usually follow contamination events.

In April 2013, an Oregon wheat farmer discovered a strange strand of wheat growing where no wheat previously had been planted. The U.S. Department of Agriculture (see USDA/APHIS 2018) ultimately identified this strand as a genetically-engineered (GE) strain developed by Monsanto and called the Roundup Ready strain (hereafter, the GE/RR wheat strain). In 1997, Monsanto began development of the GE/RR strain, and over the 1998-2005 period tested this strain on 100 field sites in 16 U.S. states, before ultimately pulling its

application. In the end, the GE/RR strain of wheat was not approved by the USDA (Jalonick, 2013; Thukral and Maeda, 2013; Boone, 2013). This un-approved GE/RR wheat strain ultimately contaminated non-GE wheat supplies not only in Oregon but nationwide. In the immediate two-month aftermath of the discovery of the GE/RR wheat strain on the noted Oregon farm, a number of news sources provided reports of the contamination's potential consequences. Klein (2013) reported alleged consequences that included South Korea and Japan (two important importers of U.S. wheat) cancelling some contracts to purchase U.S. wheat; the European Union having implored its 27 member nations to increase inspections of imported U.S. wheat; and certain wheat futures contract prices having fallen on the Chicago Board of Trade.

Given the non-approved status by the USDA of the GE/RR wheat strain, and given its contamination of various U.S. state supplies of non-GE wheat, court cases were filed in various U.S. states against Monsanto. Such cases basically alleged that farmers would incur lost sales and depressed wheat market prices due to the effects of such contamination of non-GE U.S. wheat supplies by the GE/RR strain. The filings by farmers further contended that Monsanto should have known that it was virtually impossible to prevent the GE/RR strain's contamination of other non-GE wheat supplies given that the U.S. wheat system comingles wheat deliveries from thousands of U.S. farms for sales and shipping (Jalonick, 2013; Thukral and Maeda, 2013; Bjerga, 2013; and Boone, 2013). As explained by Bjerga (2013), the actual or perceived existence of such a GE wheat strain as RR would likely dampen U.S. wheat export demand and lower U.S. wheat prices insofar as a multitude of U.S. wheat-importing clients are averse to genetically-engineered supplies of grain and other agricultural products.

In May 2013 when the GE/RR wheat strain contamination incident occurred, projections suggested that about 140 to 150 million tons of wheat were globally traded, and about a third of

this trade flow was U.S.-sourced, rendering a sizeable source of payments to U.S. wheat farmers. Japan alone was importing about 40 million tons of wheat annually and sourcing about a third from the United States. “GE-adverse” Asian wheat importers – Japan, South Korea, and the Philippines in particular -- were reported as very concerned over the May 2013 GE event, and consequently were re-considering their wheat import tenders and deals until the U.S. government assured that U.S. wheat export supplies were not contaminated with the GE product. Anecdotally U.S. Department of Agriculture (USDA/ERS 2013) data available at the time of litigation suggested that U.S exports of non-durum wheat declined 26.5% during the first few months after the May, 2013 announcement.

Purpose

In his oft-cited article concerning the use of econometric work in the courtroom, Rubinfeld (1985, p. 1048) noted that “the use of statistical methods for resolving disputes has found increasing acceptance within the adversary system,” and has “. . . opened the door to law-related econometric studies, particularly in connection with the use of multiple regression models.” The 2013 U.S. wheat contamination incident and similar episodes with other U.S. agricultural products reviewed below generally have the potential to substantially decrease U.S. farmer revenues through reduction in U.S. export sales of the relevant commodities and through lower farm prices received. The literature has provided econometric methodologies to estimate adverse GE event-induced price effects that could be tied to declines in farm revenues and income for the nation-wide crop sector subject to the contamination. Yet to date, some 34 years after Rubinfeld’s (1985) comments above, the literature does not provide a methodology to assess how such contamination events negatively influence prices and receipts at the farm-specific plaintiff level (Carter and Smith, 2007; Li et al., 2010).

Immediately after the 2013 GE/RR event, court cases were filed by farm-specific plaintiffs that produce the non-durum classes of wheat. The lack of a methodology to help these wheat farmer plaintiffs highlighted a long-standing and continuing need by the economics profession for an appropriate methodology. Our goal here is for an applied methodological contribution lacking in the literature, in particular extending econometric sector-wide studies focused on the national crop market effects to estimation of effects on farm-specific plaintiffs so as to enhance our profession's participation/usefulness in related litigation. Guided by prior relevant literature that uses time series co-integration modeling techniques to discern contamination-induced national price effects for other crops, we estimate U.S. wheat sector effects induced by the 2013 wheat contamination event, and then use these sector level effects to estimate farm-specific plaintiff effects of the contamination on farm-level prices, sales receipts and farm income for use by plaintiffs in litigation that followed the event.

Prior Related Work

Two studies that applied time-series econometric methods to estimate sector wide effects of GM variety contaminations of U.S. corn and rice supplies were deemed relevant to our study on wheat. Both studies applied time-series econometric models at the crop sector level to capture long-run reduced form co-integrated relationships among prices for a U.S. crop sector supply contaminated with a GM variety as well as for a substitute crop.

Carter and Smith (2007) used co-integration and other time series methods to examine the 2000 contamination of U.S. food corn supplies by the StarLink corn variety, a GM strain not approved by the U.S. government for human consumption. They proposed a forecasting model based on co-integration between the price of U.S. food corn and the price of sorghum (a corn substitute in certain usage categories) not directly affected by the event. They discerned a

StarLink event-induced departure from a stable co-integrating relationship among U.S. corn and sorghum prices assessed through tests of structural breaks to identify the relevant Starlink event dates. They split the entire sample into pre- and post-contamination subsamples. They then estimated the magnitude of the StarLink event effects through a comparative analysis of the difference in two sets of post-event prices: observed U.S. corn prices and what the “benchmark” U.S. food corn prices would have been had the event not occurred. Among their results, Carter and Smith (2007, pp. 531-533) concluded that the StarLink corn variety’s contamination of U.S. food corn supplies elicited a 6.8 percent decline in U.S. corn prices that persisted for at least a year.

Li et al. (2010) used time-series co-integration modeling methods to assess impacts on U.S. rice market prices from the 2006 contamination (officially announced on August 18) of U.S. rice supplies by the Liberty Link Rice 601 (LL601) variety that was not approved by the U.S. government for human consumption. They used methods of Engle and Granger (1987) and Engle and Yoo (1987) and tested for co-integration between weekly U.S. rice futures prices affected by the LL601 contamination and prices of the closest discernable substitute in global markets, namely Thai rice (Li et al. 2010, p. 32). Without having formally tested for the contamination-induced structural breaks, they assume that such contamination-induced structural breaks had occurred and then split their full sample into pre- and post-contamination subsamples. They then tested for, found, and estimated a long-run equilibrium relationship between the weekly U.S. and Thai rice prices, following the methods of Engle and Granger (1987), and used the residuals of this relationship to estimate a two-price error correction model (ECM). The ECM was then used to forecast prices over a post-LL601 period and to determine what “benchmark” prices would have been without the LL601 contamination event (Li et al. 2010, pp.

33-34). Among their results, Li et al. (2010, pp. 34-36) concluded that LL601 contamination-induced U.S. rice price declines were dramatic but short-lived, with U.S. price having declined up to about 17% within two weeks of the contamination and having fully recovered within a month of the announcement.

Stages of the Assessment Methodology for the non-GE Wheat Contamination Incident

A reliable analytical procedure for empirically assessing the impacts on relevant U.S. product markets and on firm-specific plaintiffs affected by such an incident as the May 2013 GE event would be comprised of several steps.

- First, one discerns the products allegedly affected by the monitored event, here the wheat classes affected by the May 29, 2013 incident. The four (non-durum) classes of U.S. wheat potentially affected by the event were winter white soft wheat, winter red hard wheat, winter red soft wheat and spring red hard wheat. This phase constitutes stage 1 of the proposed methodology and needs no further discussion. A plot of the monthly four U.S. wheat class prices is provided below in Figure 1, namely soft red winter wheat, hard red winter wheat, spring red hard wheat, and white winter wheat. Winter soft red wheat primarily is grown in parts of Texas, Louisiana, Mississippi, Arkansas, Tennessee, Missouri, Illinois, Indiana, Ohio, Georgia, South Carolina, North Carolina, Virginia, Maryland, Delaware, and Pennsylvania. Winter hard red winter wheat largely is produced in parts of Texas, Oklahoma, Kansas, Nebraska, New Mexico, Colorado, Wyoming, South Dakota, Montana, and California. For the most part, white winter wheat is grown in Idaho, Washington, and Oregon. Finally, hard red spring wheat is produced in Washington, Idaho, Montana, North Dakota, South Dakota, and Colorado. For considerations of space, and given our above-noted analytical purpose, we focus

throughout only on one chosen class of affected U.S. wheat, namely spring red hard wheat, while noting that our methodology developed here may be replicated for discerning the impacts on the receipts and prices of the other three U.S. wheat classes. Note too that spring red hard wheat and winter wheat are indigenous to the area of the unapproved GE/RR wheat strain. Our efforts applied to a system of four prices were guided by the bivariate co-integration modeling procedures of Carter and Smith (2007) and Li et al. (2010) that exploited the long-run co-integration relationship among the contaminated crop price and the price of the nearest discernible substitutes.

- Second, one must empirically discern the incident's impacts on national U.S. product prices, with our focus here illustratively placed on the price of U.S. spring red hard wheat (hereafter PSPRRDHD). The USDA's National Agricultural Statistical Service (NASS) publishes monthly prices for the previously mentioned four non-durum wheat classes (see USDA/NASS 2013). For reasons discussed below, this stage includes the specification, estimation and application of a state-of-the art co-integration vector auto regression or co-integrated VAR econometric model (sometimes denoted as a vector error correction or VEC model) of all four U.S. wheat class prices, including PSPRRDHD on which attention is centered here. Given Carter and Smith's (2007) useful review of an array of potential modeling strategies, we chose, for reasons cited below, from this array, the strategy that combines a reduced form co-integrated VAR model of substitute prices with a "shifter variable" defined for the specific contamination event, whereby the "shifter" is useful in discerning the event-induced price impacts.
- Third, one calculates national damages to U.S. producers of the relevant product market, here all U.S. farmers of spring red hard wheat. We develop a methodology to estimate

declines in PSPRRDHD and declines in U.S. farm receipts from sales of spring hard red wheat that are considered damages attributed to the May 2013 GE event.

- Fourth, we demonstrate how one would use econometrically-estimated national price declines attributed to the event to the firm-specific records of plaintiffs involved in the class action suit that arise from the monitored action -- *hypothetically* for a single plaintiff farmer of spring red hard wheat in response to the May 2013 GE event.

Anonymous reviewers insightfully suggested that we apply this methodology to all five U.S. wheat classes (including durum) rather than just to our chosen single class, red hard spring wheat. We opted to limit the study to this single illustrative wheat class largely based on the fact that spring red hard wheat is native to the 16-state area of the GE/RR wheat contamination. We also appeal to Sims' (1989) principle that a meaningful modeling effort should be, in part, tailored to the analytical purpose at hand. Insofar as farm-specific litigants' own financial records would have been ideal here, such records are not within the public domain. So in order to achieve our second-tier goal of translating national impacts to the farm-level, we use records for a hypothetical spring red hard wheat farmer.

The relevant studies in the extant literature discussed previously had the goal or analytical purpose of econometrically estimating the adverse national or U.S. crop sector-level price impacts of a contamination event: Carter and Smith (2007) for corn and Li et al. (2010) for rice. Our study's analytical purpose is two-tiered and is expanded to not only econometrically estimate the 2013 GE/RR wheat contamination effects on national U.S. wheat class prices (tier-1) but also to provide a methodology, lacking at the onset of the 2013 GE/RR contamination event, that is useful in litigation and that translates the national effects to Plaintiff/farm-specific declines in prices received, revenue lost, and implied declines in farm income (tier-2).

Supported by Sims' (1989) well-known modeling criteria, and given this extended analytical purpose at hand, we chose to limit the study to one illustrative wheat class in order to avoid excessive page length and to avoid what we believe would be an overly "replicative" array of four similar non-durum wheat class calculations and analyses. As noted above, we believe our analysis illustratively provided for spring red hard wheat can be seamlessly adopted by farmers of other wheat classes and to other non-wheat crop markets.

Using Co-integration "Reduced Form" Models for Contamination Events.

Carter and Smith (2007, pp. 524-526) discussed a number of alternative modeling methodological options in order to estimate adverse price effects for a U.S. crop with an emphasis on corn. Their idea is to compare, for a post-event period, observed crop prices with benchmark crop price forecasts reflecting what such post-event prices would have been had the contamination event not occurred. The "would-be" post-event benchmark prices would be generated by a structural demand and supply model of the contaminated US crop sector or market. Another of Carter and Smith's proposed options would include a "shifter" variable in the estimated U.S. structural crop sector model. The shifter's coefficient estimate could then be used to discern significant event-induced price effects (Carter and Smith, 2007, p. 526).

Yet for U.S. crop sector markets generally, Carter and Smith (2007) noted the well-known specification and enormous data procurement problems of estimating such structural models. With problems that were similarly noted for modeling the national U.S. corn market, our U.S. wheat market is comprised of four wheat class submarkets; there is a wide array of client demands for each submarket (domestic food, domestic feed, foreign food, foreign feed demands); and each demander's preference function in each submarket is unique (see Carter and Smith 2007, pp. 524-527). Carter and Smith then recommended using reduced form co-

integration modeling methods that exploit the high-quality U.S. crop price data published by NASS. These reduced form models exploit the long-run equilibrium or co-integrating relationships of the price of the contaminated crop and prices of substitutes that were not directly contaminated. Such reduced form models could avoid the noted specification and data collection problems of estimating the contaminated sector's structural demand-supply model (Carter and Smith 2007). As an alternative to their use of a reduced form co-integrated model of substitute prices combined with structural break analysis, they suggested taking the co-integrated VAR model and combining it with a "shifter" variable that may be used to discern event-specific national crop price effects (Carter and Smith 2007, p. 525).

For reasons detailed below, we follow this latter Carter and Smith (2007) strategy and combine the reduced form co-integrated vector autoregression (VAR) model (or vector error correction or VEC model) of four prices of U.S. wheat classes (well-known substitutes in consumption and production) with a "shifter" variable. Our shifter variable follows Juselius' (2006) widely applied "permanent shift dummy or binary" variable of use in empirically discerning impacts of such event-specific events as the 2013 U.S. wheat supply contamination.

We conducted Juselius' (2006, ch. 9) "known-beta" test for time-invariance of estimated parameters, with time-variance of such parameters considered as a "structural break." Since our test results (provided below) suggested that the contamination did not result in time-variance of the estimated parameters in our model, we chose not to follow Carter and Smith's (2007) strategy under conditions of a structural break and where the full sample is split into pre- and post-contamination subsamples. We rather opted for Carter and Smith's (2007) alternative of combining the co-integrated VAR model with a "shifter" variable, that is with Juselius' (2006) permanent shift dummy variable in the co-integration space. Use of such a dummy variable is

valid over the full sample when the event for which this variable is defined (here the 2013 U.S. wheat contamination) does not induce time-variance of estimated parameters or a “structural break.”

The prime focus of this analysis is to combine widely applied econometric tools with basic accounting procedures to empirically assess contamination-induced adverse negative impacts on wheat prices, and then illustratively apply the methodology to calculate plaintiff-specific effects for one of the four non-durum wheat classes, spring red hard wheat. As a result, we do not dwell on the widely-accepted and applied econometric tools themselves, but rather demonstrate how such tools are applied. Consequently, we provide only summaries of the econometric output when required of the methodology developed here.

Background on co-integration

Achieving our analytical purpose would first require estimated price impacts of the event using a highly periodic (here monthly) time-series econometric model of the four non-durum U.S. wheat class prices (see Figure 1), since they are substitutes in production and consumption. For reasons discussed below, we chose the modeling approach known as the co-integrated VAR or VEC model approach to model the four U.S. wheat class prices in order to capture the long-run forces (co-integrating relationships or CVs) that drive the individually non-stationary prices to move in a general tandem pattern.

It is well known that economic variables often fail to achieve conditions of weak stationarity required of valid inference.¹ Prior to the work by Engle and Granger (1987) on co-integration, econometric analysts often first-differenced non-stationary variables that were used in modeled estimations.²

In cases such as ours, with four individually non-stationary prices, Johansen and Juselius (1990), following Engle and Granger (1987), demonstrated that individually non-stationary series may form a group of inter-related or co-integrated variables that behave as a group in a stationary manner over time. As a result, the traditional remedy of differencing such individually non-stationary variables would jettison important long-run components (that is, co-integrating relationships or CVs) that are crucial in explaining long-run behavior. So while differencing such co-integrated variables would achieve stationarity of the modeled time series, estimates from such differenced data would likely incur misspecification bias from omitted relevant information, the CVs. It is well-known that the co-integrated VAR approach permits the retention of levels-based CVs in a reduced “error correction” space, so as to achieve stationarity requirements (and non-compromised inference) while avoiding difference-induced misspecification bias (Engle and Granger, 1987; Johansen and Juselius, 1990).

We first achieve a statistically adequate underlying (unrestricted) model. Second, we test for and impose co-integration restrictions of the four U.S. wheat class prices. We then focus on and interpret the co-integration parameter estimates necessary to formulate a methodology that estimates GE event-induced declines in prices and receipts at the national and plaintiff-specific levels for an illustrative class of U.S. wheat, namely spring red hard wheat.

Statistically adequate underlying model and Haavelmo’s ceteris paribus conditions

It is well-known that an unrestricted vector autoregression or VAR model of the noted wheat prices in (here logged) levels posits each U.S. wheat class price as a function of k lags (here $k=6$) lags of itself and of $k=6$ lags of the three remaining prices.³ Johansen and Juselius (1990) demonstrated that such a levels VAR model may be re-written into an algebraically equivalent unrestricted vector error-correction or VEC model:

$$(1) \quad \Delta x(t) = \Gamma(1) * \Delta x(t-1) + \dots + \Gamma(k-1) * \Delta x(t-k+1) + \Pi * x(t-1) + \Phi D(t) + \varepsilon(t).$$

Supported by diagnostic evidence reported below, we specified the model with linear trends in the variables and the co-integrated relationships. The number of endogenous variables, denoted as p , is four. The $\varepsilon(t)$ are white noise residuals, the delta is the difference operator, while the $x(t)$ and $x(t-1)$ are p by 1 vectors of the endogenous variables in current and lagged levels. The $\Gamma(1), \dots, \Gamma(k-1)$ terms are p by p matrices of short-run regression coefficients, and Π is a p by p long-run error correction term to account for endogenous levels. The other terms are collectively referred to as the model's short-run/deterministic or "short-run" component. The $\Phi D(t)$ corresponds to a set of deterministic variables, including an array of binary variables that will be added to address stationarity and specification issues discussed below. From Engle and Granger (1987) and Johanson and Juselius (1990), the error correction (EC) term is decomposed as follows:

$$(2) \quad \Pi = \alpha * \beta'.$$

The α is a p by r matrix of adjustment coefficients or "alphas" where r is the number of co-integrating relationships or the rank of Π discussed below. The β is a p by r vector of co-integrating parameters or "betas." Π captures level-based long-run information: stationary linear combinations of the non-stationary levels data under co-integration, permanent shift binaries to capture long-run impacts of policies and/or market events, and a linear trend.

We followed the recommended methods of Juselius (2006) to include several sorts of binary variables to address an array of stationarity and specification issues. A set of eleven binary variables was included to account for seasonality of U.S. wheat crop prices. As well, we included the following "permanent shift" binary variables in levels within Π and in differenced form within the short-run component of equation (1):

- RECESS in order to account for the wheat market effects of the U.S. economic recession during December 2007 through June 2009.
- FB2008 in order to account for the U.S. wheat market impacts of the 2008 U.S. farm bill.
- GEEVENT in order to capture the immediate influences on the four endogenous wheat class prices of such an event as the above GE event that occurred in May 2013. Valued at unity for June 2013 and July 2013, and at zero otherwise, this binary variable accounts for the event-induced market effects such as cancellation of foreign import orders for U.S. wheat, coincidental wheat futures market price declines, among other things, reported during the immediate months following the event's May 2013 occurrence by Bjerga (2013), Boone (2013), Klein (2013), Jalonick (2013), and Thukral and Maeda (2013). Although the U.S. wheat contamination was discovered in April, we opted to not to include April in GEEVENT's definition because we located no evidence of a news leak before the USDA's official announcement on May 29, 2013. Since the USDA's announcement of the GEEVENT event was on May 29, 2013, with most of the month (all but two full days) having already passed, and with little or no possible GE/RR event influence on the USDA's estimates of the May 2013 wheat class prices, we chose to begin GEEVENT with June 2013 (see USDA/APHIS 2018) and exclude May 2013 that some may have preferred to have been included. This GEEVENT binary is of particular use in our study because in addition to being consistent with the national contamination-induced effect estimates in tier-1, the binary's definition is also consistent with tier-2 effect estimates since all farmer Plaintiffs who filed in the actual court case did so for marketings of spring red hard wheat sold over the June-July 2013 period.

Moreover, following Juselius (2006), other “outlier” binary variables were included in equation 1’s short-run component. These binary variables were designed to account for “extraordinary” and exogenous wheat market influences attributed to short-run events not of direct relevance to the model being estimated.⁴ Juselius (2006) recommends this binary variable-based procedure that focuses on achievement of Haavelmo’s (1944) well-known conditions that use “statistical devices” that clear the data of extraneous exogenous effects not of direct theoretical relevance to the modeled market relationships. These binary variables are included only in the model’s short-run component and not in the long-run cointegration space. Consequently we follow Clements, Tidwell, and Jin (2017) and, aside from indicating from which exogenous events they may arise, do not analyze or discuss these binary variables at length.

Following Juselius (2006), when a binary variable was identified, the variable or variable group was included in the appropriate parts of equation 1, equation 1 was re-estimated, and the binary variable was retained if a battery of diagnostics suggested enhanced specification. We do not focus directly on the econometric output that is not directly related to our analytical goal of developing a methodology to assess damages of such an event as the May 2013 GE/RR event. However, as exhibited in Table 1 we demonstrate that we achieved a statistically adequate underlying model.

Table 1 permits a comparative analysis of evidence from a battery of diagnostics generated by the first and last estimations of an unrestricted VEC (equation 1) in our quest for enhanced specification. Prior to specification enhancement efforts, evidence suggested issues with non-normal residual behavior, particularly at the univariate level for all four endogenous prices. After a series of re-estimations with statistically supported binary variables discussed

previously, the underlying unrestricted model saw the trace correlation, a goodness-of-fit indicator, nearly double to 0.626; evidence suggested that serial correlation and heteroscedasticity were not issues; and indicators of skewness and kurtosis improved and ultimately fell within accepted levels. From this information, we conclude that we met literature-established evidential standards that permit us to conclude that we achieved a statistically adequate underlying unrestricted model for which co-integration properties may be tested and exploited. The diagnostic results exhibited in Table 1 not only render us confident that inclusion of the various binary variables noted previously was accomplished in a properly implemented procedure as described in Juselius (2006) but also that we avoided well-known econometric consequences that potentially arise from excessive use and/or misspecification of such binary variables.

Testing for co-integration, imposing appropriate rank of Π , and testing for time-invariance of parameter estimates.

Equation (2) shows that Π is a $p \times p$ (here 4 by 4) matrix equal to the product of two p by r matrices, where r is the reduced rank of Π and represents the number of co-integrating relations or CVs that error-correct the system of our four individually non-stationary prices. Testing for the appropriate r is a test for co-integration. Under co-integration, the rank of $\beta'x(t-1)$ is reduced from $p=4$ (i.e., $r < p$), and r may range from 1 to 3. The literature has widely employed Johansen and Juselius' (1990) nested trace tests to ascertain Π 's reduced rank or r . Table 2 provides nested trace test statistics and test results. Evidence at the 5% significance level was sufficient to reject the first null hypothesis that $r \leq 0$ and was insufficient to reject the second (that $r \leq 2$) as well as the third and fourth null hypotheses. These nested results suggest that Π 's reduced rank is $r = 1$ and that a single co-integrating vector (CV) corrects the system. A reduced rank of $r = 1$ was then imposed on Π .

As instructed by Juselius (2006, ch. 9) one should test for the time-invariance (or constancy) of estimated parameters, and she offers the “known-beta” test programmed in Estima (2013) to do so. The known-beta procedure (see Juselius 2006, ch. 9) formally tests the null hypothesis that parameter estimates of the co-integrated model (with the imposed restriction of Π having the rank of here $r=1$), including GEEVENT defined for the 2013 U.S. wheat contamination, are constant or time-invariant over the entire sample including the sample’s June-December, 2013 sub-period after the time of the contamination’s occurrence. Following Juselius (2006, ch. 9), and as well Babula et. al’s (2006, p. 47) application of the test to a monthly co-integrated VAR analysis of monthly prices of U.S. sugar-related products, test values (normalized by a 5-percent critical value and asymptotically distributed as a chi-squared distribution) were all less than unity. This finding in turn suggests that the parameter estimates of our model (including the GEEVENT estimate) are time-invariant, and hence are valid over the entire sample period. Evidence and analysis below suggests that while the contamination event modeled by GEEVENT did have an effect, the contamination event did not elicit time variance of coefficients (did not elicit a structural break), and the estimates imputed from GEEVENT are valid. In other words, formal statistical evidence suggests that for the 2013 wheat contamination event a structural break or parameter time-variance was not an issue.

This statistical evidence that the 2013 wheat contamination of focus did not elicit a structural break is the principal difference between this study and the Carter and Smith (2007) and Li et al. (2010) enquiries on the US price effects of the US corn and rice contaminations, respectively. Ours is the only one of the three studies where formal statistical analysis (known-beta test results above) suggested that there was not a structural break. So given the time-invariance and valid interpretation of the shifter or the binary GEEVENT’s coefficient over the

whole sample, one is not compelled to split the full sample, as did Carter and Smith (2007) and Li et. al. (2010), into pre- and post-contamination subsamples so as to ultimately estimate contamination-induced price effects.

The time-invariant GEEVENT coefficient is valid over the full sample (before and after the 2013 contamination event), and being defined for the contamination event per se, the GEEVENT may be used as discussed below to discern empirically the contamination-induced price effects (Patterson 2000, p. 195; Juselius 2006). In our study's case without a contamination-induced structural break, and under the important caveat that our four modeled U.S. wheat price variables provide a statistically adequate reduced form cointegration model as our specification discussion above shows that we have done, then one should not split up the sample into pre- and post-contamination subsamples. And as well, the importance of finding a non-wheat or non-US wheat substitute price as a non-contaminated "control" as Carter and Smith (2007) found for corn in sorghum price is not as crucial or necessary. Given a statistically adequate cointegrated VAR model of the four substitute wheat prices, and given our finding (Table 2) of one cointegrated relationship normalized on spring red hard price, then the GEEVENT likely ferrets-out the contamination-induced effects on US spring red hard wheat price without needing to throw-in a control variable from an outside market. Finally, Carter and Smith's (2007) recommended modeling strategy of using a cointegrated VAR or VEC model with a "shifter" must have been for cases such as ours without structural breaks over the entire sample; otherwise the recommended "shifter's" coefficient estimate would not have been constant (that is would have been time-variant) and not valid for discerning contamination-induced effects.

An estimate of GE event-induced price declines on spring red hard wheat

Following Johansen and Juselius (1990) and Juselius (2006), we normalized the co-integrating relationship on the price of spring red hard wheat (PSPRRDHD), and subsequently conducted a series of hypothesis tests concerning model specification. We imposed the restrictions that were accepted by the hypothesis tests, and then re-estimated the single CV normalized on PSPRRDHD.⁵ The coefficient on the GEEVENT binary variable was the part of the CV of direct relevance to our previously discussed analytical purpose.

When the CV is normalized on PSPRRDHD, and since the estimations were done in double-log format, then one may invoke Halvorsen and Palmquist's (1980) well-known convention of interpreting the GEEVENT coefficient estimate of -0.03909159. Following this convention, an estimate arises of -3.83 percent for the GE event-induced reduction in the price of spring red hard wheat. More specifically, we raised e , the base of the natural logarithm, to the power of the econometric GE-event binary coefficient estimate (-0.03909159), subtracted 1, and multiplied by 100 to get -0.03833787. The coefficient suggests that on average, the price of spring red hard wheat was about 3.83% lower during the two-months following occurrence of the May 2013 GE event (the GEEVENT binary's period of definition) than during other parts of the sample period. This reduction in price is consistent with findings of a contamination-elicited price decline for the commodities of focus in Carter and Smith (2007) for corn and in Li et al. (2010) for rice.

Estimated national effects on price and receipts from the May 2013 GE event

Table 3 provides a method to calculate the national effects attributed to the May 2013 GE event on price and receipts of U.S. spring red hard wheat farmers. These effects demonstrate the calculation procedures employed in late-2013/early-2014 when actual court cases were filed in

response to the noted event. As such, the procedures use the most reasonable assumptions and the most accurate projections of price reductions, spring red hard wheat production, and crop percentages sold on the spot market as were available by economists and attorneys who participated in the noted class action cases.

Column A provides the projected price of spring red hard wheat, PSPRRDHD, the average of the USDA's monthly PSPRRDHD values for the June-December, 2013 period immediately following the May, 2013 GE event (USDA/NASS, 2018).⁶ Column B provides the USDA's projected June 2013 – May 2014 spring red hard wheat production (490.39 million bushels) as reported in the *Wheat Outlook* report that had most recently been published in the immediate aftermath of the May 2013 event and available to litigators in late 2013 (see USDA/ERS, 2013). The impact estimates on national spring red hard wheat receipts are therefore conditional estimates and hinge on assumed proportions of the post-event June, 2013/May, 2014 market year crop that would have been sold on the spot market immediately after event. Generally unknown, we provide four alternative assumptions for these proportions in column C.

Column E provides projected declines in U.S. receipts for spring red hard wheat attributed to the May, 2013 GE event. The econometrically-estimated event-induced decline in PSPRRDHD, -3.83% is in column E. As a result, column D provides the projected damages, in terms of decreased receipts to U.S. spring red hard wheat under four alternative assumptions of percentages of the crop sold onto the spot market in late-2013/early-2014.⁷ Column F is the estimated 3.83 percent GE event-induced price reduction estimate applied to \$6.97 projected price in dollars per bushel.

In summary, using our methodology as exhibited in Table 3 would generate estimated damages from the May 2013 GE event to U.S. spring red hard wheat farmers of lost receipts that ranged from \$32.76 million to \$131.06 million from a GE event-induced drop in PSPRRDHD of 3.83% or \$0.27 per bushel. Because of space considerations, and given our goal of methodology development, we provide such estimates in Table 3 only for spring red hard wheat, and remind readers that analogous tables may be calculated for the other three noted U.S. wheat classes.

Estimated effects on a hypothetical plaintiff farmer's price and receipts from the May 2013 GE event

Table 4 provides how the econometrically estimated drop in price, -3.83%, may be used to estimate plaintiff-farm-specific declines in spring red hard wheat price and in receipts from the May 2013 GE event. Given the confidentiality of farm-specific financial records that may not be published herein, the data exhibited in Table 4 are based on records of a *hypothetical* plaintiff farmer's wheat transactions during June 2013, soon after the event. In conferring with extension economists who possess actual records, these hypothetical records were deemed reasonable. The methodology suggests that occurrence of the May 2013 GE event resulted in a reduction of \$0.31 per bushel price received from \$8.04 to \$7.73. The actual price received of \$7.73 per bushel (column 3) is adjusted into a higher (hypothetical) price that would have occurred had the GE event of May 2013 not taken place -- \$8.04 per bushel (column 4). This latter hypothetical price of \$8.04 per bushel was obtained by dividing the econometrically-estimated price decline factor of 0.961662626 that in term is the base of the natural logarithm, 3, raised to the power of the previously noted GEEVENT coefficient estimate (-0.03909159) and subtracting 1.0.⁸

Further, this paper provides a contribution by taking national price effect estimate from modeling strategies similar to Carter and Smith's (2007) and Li et. al. (2010) general modeling strategies, and then translating them, for the first time in the literature, to farm-specific plaintiffs

for use in litigation. Actual and projected receipts are the volume sold multiplied by the actual and projected prices received in columns 5 and 6, respectively. In turn, the negative difference in the actual and projected receipts represents the damages in terms of lost farm receipts attributed to the May 2013 GE event in column 7. In this *hypothetical* plaintiff's case, our methodology estimates that \$4,807 in lost receipts was attributed to the May 2013 GE event.

As insightfully noted by an anonymous reviewer, there are other factors not accounted for which could increase or decrease the adverse economic effects of the 2013 GE/RR contamination of US wheat supplies. These other factors include various transaction costs associated with "cleaning" the wheat in the U.S. supply chain, costs of testing for the presence of a GM contaminant, and restructuring of export shipments from countries who imposed the ban and those countries that did not.

Summary and Conclusions

In May 2013, the U.S. government annouced a genetically-engineered strain of wheat contaminated non-genetically engineered supplies of U.S. wheat. As a result, and within two months of the event, an array of consequences emerged that included, among other things, cancellation of a number of foreign nations' purchase order contracts of U.S. wheat and declines in wheat prices and wheat futures prices. It was widely reported that such alleged consequences associated with the May 2013 GE event would likely result in declines in U.S. wheat prices and U.S. wheat receipts (Boone, 2013; Bjerga, 2013; Jalonick, 2013; Klein, 2013; and Thukral and Maega, 2013). Class action by certain wheat farmers against the developer of the contaminating GE wheat strand soon occurred in late-2013/early-2014.

Previous studies utilized time-series co-integration modeling methodologies to assess U.S. crop sector-level price declines induced by crop contamination events (Carter and Smith,

2007; Li et al., 2010). Yet to date, no studies have been located that extend the sector-level adverse price effect estimates to the farm-specific plaintiffs for use in litigation that often ensues such contamination events and that ensued 2013 GE/RR wheat contamination incident. Consequently, we develop a methodology with which to accurately empirically assess the damaging impacts of the May 2013 GE event on U.S. wheat farmers nationally and at the plaintiff farm-level. Implicit in this specific case was a need in establishing a methodology by which damages from such a crop supply contamination event may be assessed for producers generally at the national and at the producer/plaintiff level.

This methodology combines econometric modeling and simple accounting procedures in order to empirically assess national and plaintiff farm-specific damages in the form of declines in wheat price and wheat receipts that were estimated to have emerged from contamination of the U.S. wheat supply in May 2013. Throughout we illustratively applied the methodology to the national market for and to a *hypothetical* individual plaintiff farmer of the U.S. spring red hard wheat class and within the time frames required of the class action litigation that occurred after the actual May 2013 GE event. We offer this summary of the methodology as a more general way for the economics profession to similarly assess damages for contamination of other U.S. wheat class markets and more generally for other commodity or food market supplies that may become contaminated.

The methodology entails four steps: (1) to determine the product market affected; (2) to develop a method to model appropriate markets in order to empirically assess the monitored event-induced effect on the product price; (3) to apply the estimated event-induced price decline to production projections at the time of litigation; these projections should be the most recent available at the time of the litigation and are applied to provide estimates of declines in U.S.

receipts of the monitored market; and (4) to apply the event-induced price effect to plaintiff-specific sales records, prices received and receipts to estimate event-induced damages in terms of prices received and lost sales receipts.

The methodology was applied to the U.S. spring red hard wheat industry in the wake of the May 2013 GE event. Results suggest that U.S. spring red hard wheat farmers lost receipts that ranged from \$32.77 million to \$131.06 million, and incurred a drop in spring red hard wheat prices of 3.83% or \$0.27 per bushel. At the farm level, our methodology suggests that a *hypothetical* farmer received \$0.31 less per bushel because of the May 2013 GE event in this event's immediate 2013 aftermath. In this plaintiff's case, our methodology estimates that \$4,807 in lost receipts was attributed to the May 2013 GE event.

End Notes

¹ Throughout, non-stationary data refer to those that are integrated of order 1, $[I(1)]$, while stationary data refer to those that are integrated of order zero $[I(0)]$. Throughout, “levels variables” are modeled in natural logarithms.

² As a preliminary analysis, we performed augmented Dickey Fuller (ADF) tests on the four logged prices (see Dickey and Fuller (1981); Hamilton (1994)). Since the ADF t_{μ} values were all negative and 2.12 or less in absolute value, and well below the 5% absolute value of 2.89, evidence suggested that evidence was insufficient to reject the null hypothesis of non-stationarity in all four cases. Additionally, we employed the systems-based and rank-dependent multivariate unit root tests that emerged from the co-integrated VAR model ultimately estimated below and recommended by Johansen and Juselius (1990), Juselius (2006), Juselius and Toro (2005), and Juselius and Franchi (2007). These tests generated the following test statistic p-values that were all below 5% and suggested that evidence was sufficient to reject of the null hypothesis of stationarity: price of spring red hard wheat or PSPRRDHD ($p = 0.007$), price of winter red hard wheat or PWINRDHD ($p = 0.01$), price of winter red soft wheat or PWINRDSF ($p = 0.01$), and price of winter white soft wheat or PWINWHSF ($p = 0.04$).

³ A $k = 6$ lag structure was used by employing a likelihood ratio test lag search procedure as recommended by Juselius (2006).

⁴ Following Juselius (2016), an event that occurred over short run subsamples was considered potentially extraordinary if its standardized residual exceeded 3.0 in absolute value. We

followed recent procedures applied by Clements, Tidwell, and Jin (2017) and designed this rule on the effective sample size of T observations (126 observations in our analysis) using the Bonferoni criterion programed in Estima (RATS package, Version 8.2): $INVNORMAL(1-1.025)^T$. $INVNORMAL$ is a function for the normal distribution that returns the variable for the cumulative density function as a standard normal distribution and suggested that the absolute value of the Bonferoni value was 3.54. And following Clements, Tidwell, and Jin (2017), having seen that there were some short-term events with potentially extraordinary effects with absolute standardized residuals that fell between 3.0 and 3.54, we chose a more conservative absolute value criterion of 3.0 rather than 3.54. Four short-run binary variables were specified and are herein identified by the defining event dates for which they take unity values (zero otherwise), and were then included in differenced form in equation 1's short-run component. The first three were similar to and similarly rationalized for model inclusion in Babula's (2011) co-integrated VAR modeling study of monthly US wheat-based prices. The first two binaries were included to collectively account for the effects of the 2008 US farm bill's implementation and for the 2007-2008 global spike in commodity demands and prices: one takes a 1.0 value for the February-July, 2008 period and the other for the September, 2007-April, 2008 period. The third was defined for July 2010 through the sample's end to account for the effects of accelerating US economic activity and marked fall in global commodity prices (Babula, 2011). A fourth and final binary variable was defined for the July-August 2012 period and attempts to capture some of the 2012 anticipatory effects as the U.S. Congress moved to formulate a one-year extension of the 2008 farm bill, legislation that was implemented on January 1, 2013.

⁵ The entire co-integrating relationship is as follows:

$$\text{PSPRRDHD} = 1.198 * \text{PWINRDHD} + 1.1804 * \text{PWINRDSF} - 1.096 * \text{PWINWHSF} -$$

$0.0391 * \text{GEEVENT} - 0.202 * \text{FB2008} + 0.152 * \text{RECESS}$. In the above relation, and following earlier work on monthly U.S. wheat prices (Babula 2011, pp. 112-123), the trend is restricted to zero. Two of the four prices generated positive coefficients and we conjecture that these opposing price patterns may have arisen from tradeoffs in wheat usage and producers for whom certain classes of wheat competing for acreage. The GEEVENT coefficient -0.0391, actually generated as -0.03909159 with the latter used for damage calculations. Focused on methodology here, we realize that some may think that the model estimated should have included other variables. We relegate this issue to future research and note that in spite of what some may consider an overly simple all-price endogenous variable setting, our efforts are nonetheless valid for two reasons. First, we meet, for our chosen model of U.S. non-durum wheat prices, accepted statistical and diagnostic standards clearly reflected in Table 1. Second, our “. . . inclusion of co-integrating relationships enhances structure in that they remain invariant to expansions in the information set” (Bardsen, Fischer, and Nymoen ((1998), p. 506). So future research could expand our modeled information set while our results here remain valid.

⁶ This average price of \$6.97 is the mean of the following June-December, 2013 PSPRDHD values, published by the USDA/NASS (2018): \$7.73, \$7.29, \$6.98, \$6.72, \$6.85, \$6.70, and \$6.53, respectively.

⁷ For example, the high level of estimated lost receipts is \$131.04 million under the extreme assumption that 100% of production was sold on the spot market. This level is calculated as the

projected price of \$6.97 (A) times projected 2013/2014 market year production of 490.39 million bushels (column B) times (-0.038337374), the estimated GE event-induced reduction in PSPRRDHD. As previously noted, the latter estimate is obtained by raising e, the base of the natural logarithm, to the power of (-0.03909159), the econometric coefficient estimate generated by the GEEVENT binary variable and subtracting 1.0.

⁸ The GEEVENT binary's beta estimate is -0.03909159. The price decline factor is $[e^{-0.03909159}-1] = 0.961662626$. The actual per bushel price of \$7.73 is adjusted by the factor of 0.961662626 to render a per-bushel projected price of \$8.04 that would have occurred had the May, 2013 GE event not occurred.

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Table 1. Mis-specification tests for the Unrestricted Model Before and After Specification Enhancement Efforts

Test	Null hypothesis &/or test explanation	Prior specification efforts	After specification efforts
Trace correlation	Goodness of fit; large proportion desirable	0.345	0.626
LM test serial correlation (at lag 1)	Ho: No serial correlation of the residuals. Reject the null hypothesis when the test statistic $> \chi$ -squared critical value of 32.0	11.63	17.11
ARCH test (at lag 1)	Ho: No heteroscedasticity of the residuals. Reject the null hypothesis when the test statistic $> \chi$ -squared critical value of 135.80	105.26	104.76
Doornik-Hansen test, system-wide	Ho: The residuals of the system behave normally. Reject the null hypothesis when the test statistic $> \chi$ -squared critical value of 20.09	62.63	9.07
Univariate Doornik-Hansen tests	Ho: The residuals of each equation behave normally. Reject the null hypothesis when H-D test statistics > 9.2		
Δ PSPRRDHD		34.380	0.180
Δ PWINRDHD		13.110	0.880
Δ PWINRDSF		13.590	0.024
Δ PWINWHSF		16.650	3.630
Skewness (kurtosis) values.	Skewness: small values acceptable. Kurtosis: acceptable range < 5.0		
Δ PSPRRDHD		1.20 (8.54)	0.07 (2.92)
Δ PWINRDHD		0.75 (3.59)	-0.13 (3.13)
Δ PWINRDDF		0.81 (3.95)	0.026 (2.78)
Δ PWINWHSF		0.98 (5.12)	-0.17 (3.60)

Note: Kurtosis values are in parentheses.

Table 2. Nested Trace Test Statistics and Test Results

Null hypothesis	Trace value	Critical Value	Test Result
Rank or $r \leq 0$	86.57	63.66	Reject the null that $r \leq 0$
Rank or $r \leq 1$	38.28	42.77	Fail to reject the null that $r \leq 1$
Rank or $r \leq 2$	17.23	25.73	Fail to reject the null that $r \leq 2$
Rank or $r \leq 3$	4.85	12.45	Fail to reject the null that $r \leq 3$

Note: At the 0.05 level of significance

Source: CATS2 routines developed by Dennis (2006) in RATS (Version 8.2) of Estima.

Table 3. Estimates of National Damages to U.S. Spring Red Hard Wheat Farmers

A	B	C	D	E	F
Projected price (\$/bushel)	Projected production 6/2013-5/2014 (million bushels)	% Projected production sold on spot market	Projected lost receipts for U.S. wheat farmers (million \$)	% drop in PSPRRDHD	Drop in PSPRRDHD (\$/bushel)
\$6.97	490.39	100	-\$131.04	-3.83%	-\$0.27
		75	-\$98.28		
		50	-\$65.53		
		25	-\$32.76		

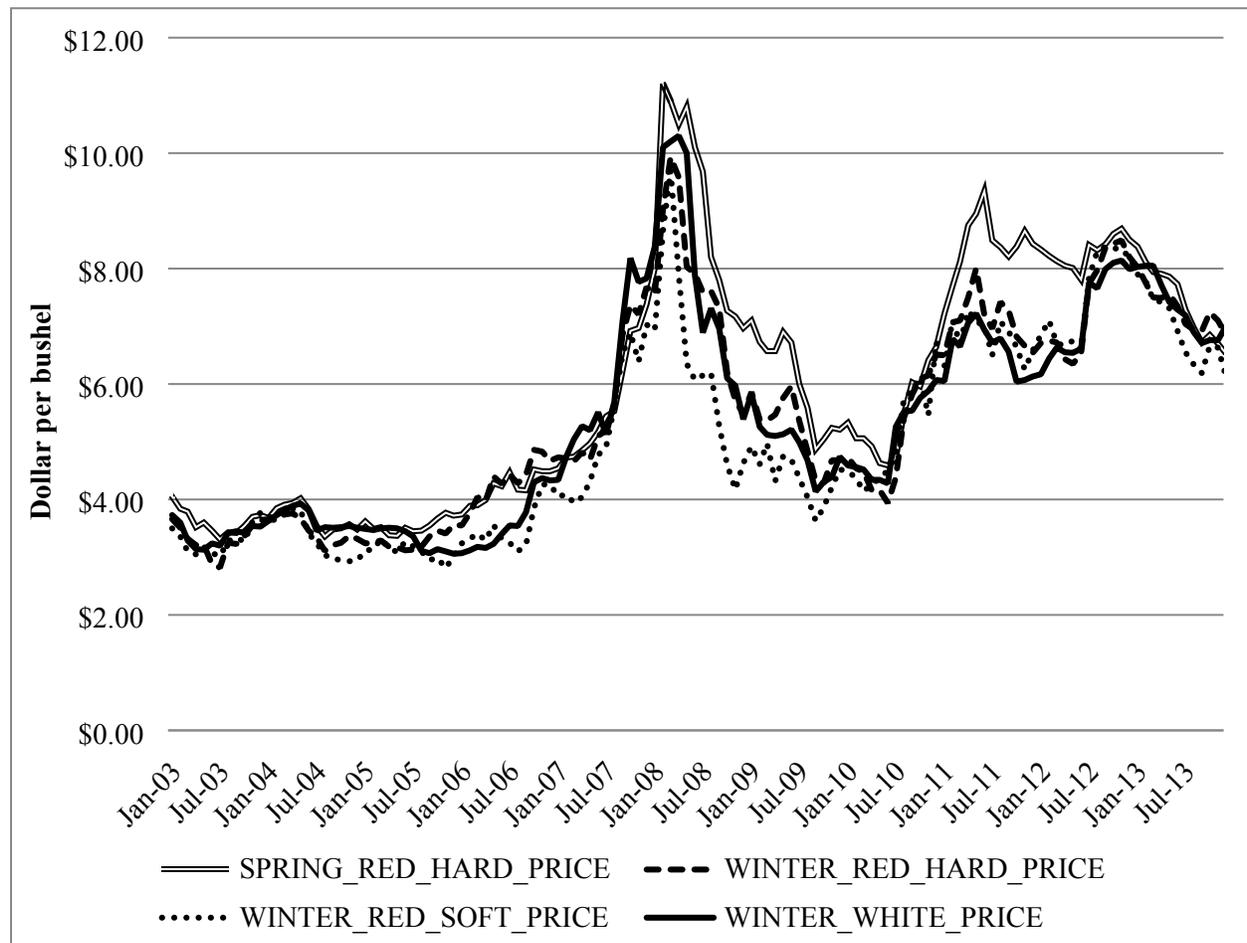
Sources. Columns A and B are based on USDA projections that were available during the noted litigation in late-2013. (USDA/ERS 2013). Column C is based on assumptions made by the authors. Column D is based on calculations using columns A, B, and C, and the price reduction estimate. The estimated coefficient on the GEEVENT binary is -0.03909159. When the base of the natural logarithm is raised to the power of this coefficient estimate, and 1.0 is subtracted from it, and this result is multiplied by 100, one obtains the econometric price reduction estimate of -3.8337374% or -0.383% reported above. So consider the -\$98.28 million loss of receipts in column D under the assumption that 75% of the U.S. spring red hard wheat was marketed on the spot market. This estimate of lost receipts is calculated as the \$6.97 price times -0.038337374 times 490.39 million bushels times 0.75. Column F is calculated as \$6.97 times -0.38337374.

Table 4. Estimation of Hypothetical Damages to a Spring Red Hard Wheat Farmer

Date	Volume sold, bushels	Actual price received	Projected price received	Actual receipts, dollars	Projected receipts, dollars	Lost receipts, dollars
June 10, 2013	15,600	\$7.73	\$8.04	\$120,588	\$125,395	\$4,807

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Figure 1: Monthly National Wheat Prices Received by Farmers, January 2003 to December 2013



Source: National Agricultural Statistical Service (NASS)