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The Impact of Renewable Fuel Production on Soybean Oil Spatial Price Dynamics

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We analyze the price relationship of crude and refined-bleached-deodorized (RBD) soybean oil prices among four regional U.S. markets (Central Illinois, U.S. Gulf, West Coast, and East Coast). Econometric time-series methods were used to detect price integration, linkages, and responsiveness for each oil type and among each market. Results show that the four markets have remained price-integrated in the long run. This implies that the markets are spatially efficient. The results, however, also suggest that this level of market efficiency may have decreased to some extent after the U.S. biodiesel production surge in the mid-2000s.

Key words: biodiesel, market integration, soybean oil, spatial price analysis, VAR, vector error correction

Vegetable oils are used for food, cooking, biofuel production, and industrial purposes. These oils are derived from processing oilseeds chemically or using a press. The result is typically a commodity oil that can be blended with other oils, or used independently of other oils. As a commodity, arbitrage of such oils occurs frequently. Oils may move spatially or be stored for short time periods. The United States is a large vegetable oils consumer, and soybean oil represents a considerable share of domestic vegetable oil consumption. Soybean oil is also one of the most widely consumed cooking oils. Of the total U.S. domestic edible oils consumption, soybean oil accounted for the largest share, 56%, in 2011 (U.S. Department of Agriculture (USDA, 2013a).

One factor that has altered the use, and subsequent geographical flow, of soybean oil is the introduction of vegetable oil-based renewable diesel fuel, which is more commonly termed *biodiesel*. U.S. biodiesel production increased dramatically from 8.6 million gallons in 2001 to approximately 967.4 million gallons in 2011 (U.S. Energy Information Administration (EIA, 2013a). Soybean oil has been the largest biodiesel feedstock. To produce biodiesel, more soybean oil has been demanded. In 2009, for example, soybean oil directed toward biodiesel use increased soybean oil consumption by 1,974 million pounds, or 0.895 million metric tons. Soybean oil used to produce biodiesel totaled 4,153

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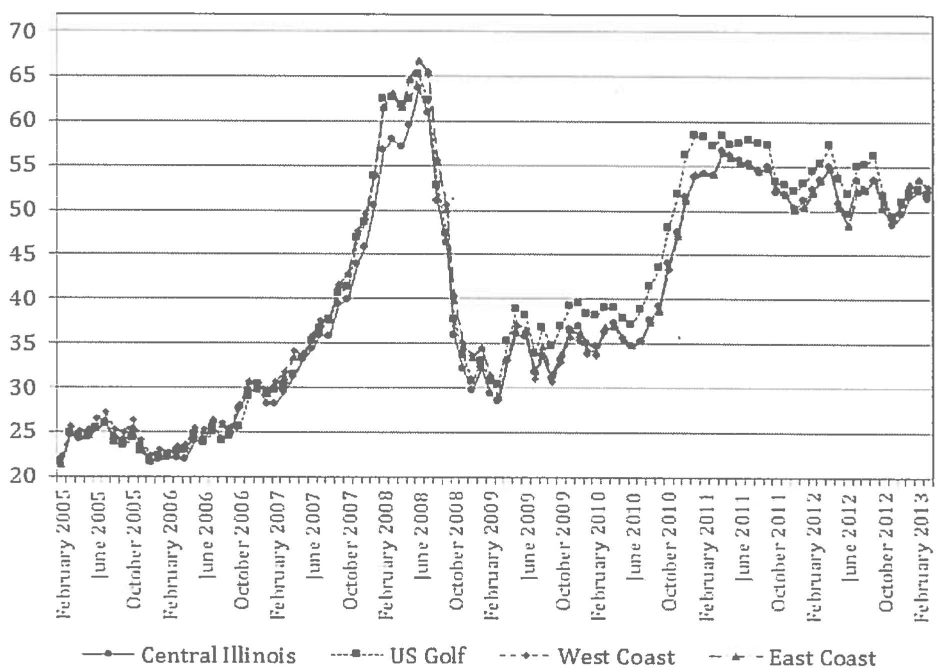
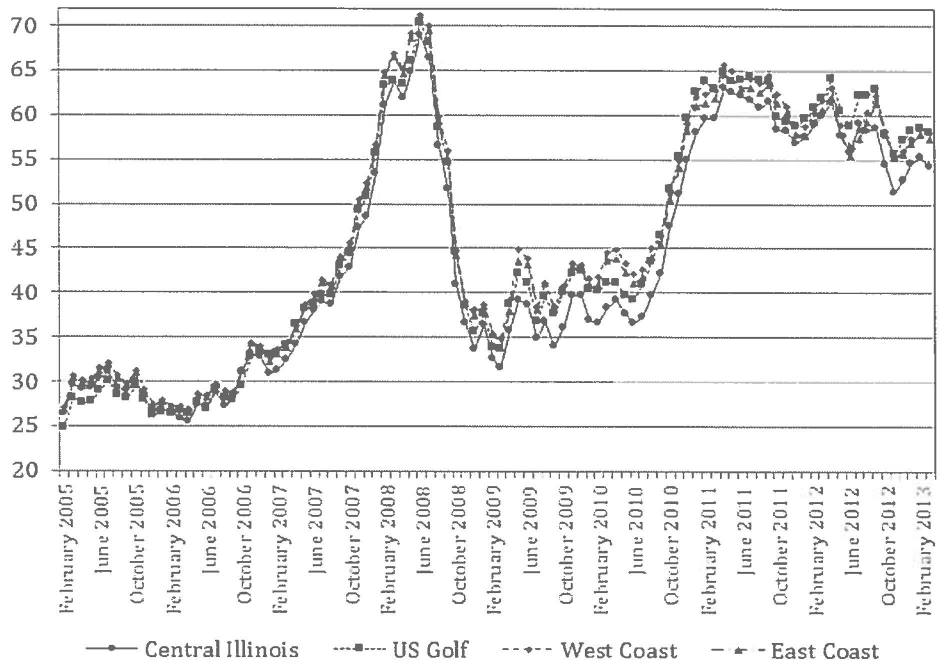
million pounds in 2011, or 1.88 million metric tons (EIA, 2013b). Biodiesel's introduction and use may represent a structural change and may have altered U.S. soybean oil spatial price dynamics.

Figures 1 and 2 display historical price trends for crude and refined-bleached-deodorized (RBD) soybean oils, respectively. These figures indicate that changes in supply-demand factors have caused significant price increases, especially in 2007 and 2010. The 2007 increase is of particular interest because it coincided with the global food price crisis and the increase in production of biodiesel. From 2005 to 2008, wheat prices increased by 127%, rice prices increased by 170% and corn prices almost tripled (Mittal, 2009). Prices of soybeans, in particular, rose by 107% between 2006 and 2008 (Steinberg, 2008). U.S. biodiesel production also experienced a sudden surge during this time—from an average of 7 million gallons in 2005 to an average of 40 million gallons in 2007 (Figure 3). These two events were also closely related. The International Food Policy Research Institute calculated that 30% of the increase in average grain prices between 2000 and 2007 was accounted for by biofuel production (Braun, 2008).



Source: The Jacobsen

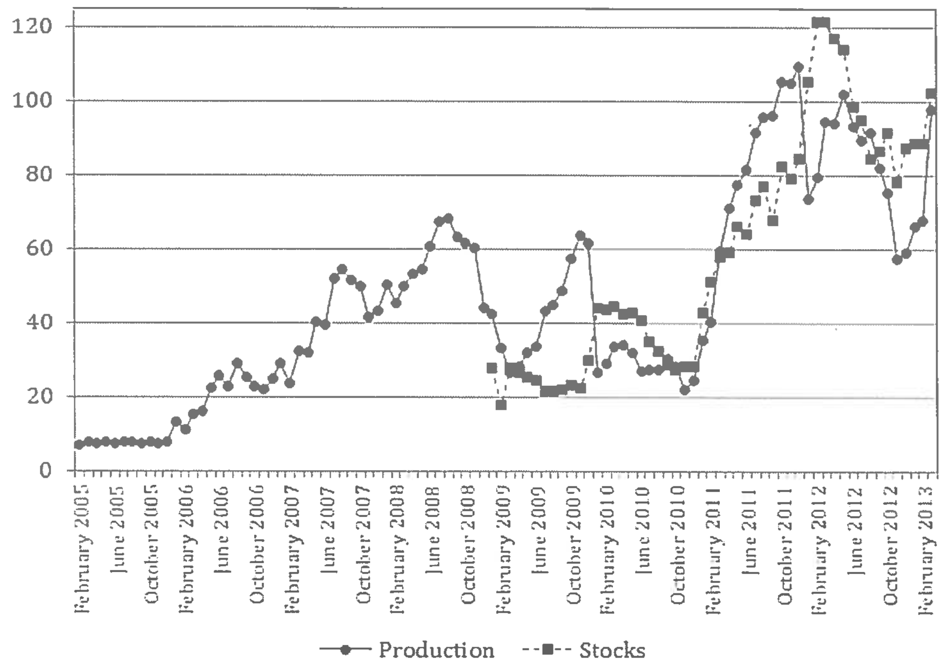
Figure 1. Monthly Average Crude Soybean Oil Prices (cents per pound) from February 2005 to March 2013.



Source: The Jacobsen

Figure 2. Monthly Average Refined-Bleached-Deodorized Soybean Oil Prices (cents per pound) from February 2005 to March 2013.

While prices of different commodities were moving in the same upward direction, these changes in prices experienced considerable variability. Of particular interest are the between-series spatial price spreads of soybean oils from different markets, which are especially notable after 2007. Although the spreads may appear minimal, even small deviations can signal considerable spatial price arbitrage opportunities. U.S. Gulf crude soybean oil prices, for example, have been consistently higher since 2009, while Central Illinois RBD soybean oil prices have been consistently lower during the same period. Because of this and given the significance of soybean oil in the U.S. economy, a more thorough examination of spatial pricing patterns for U.S. crude and RBD soybean oil is warranted. Understanding the extent to which prices of spatially separate soybean oil markets are integrated and how price relationships may have changed are crucial to price discovery and derived demand estimation.



Source: Monthly Energy Review, U.S. Energy Information Administration

Figure 3. Monthly U.S. Biodiesel Production and Stocks (million gallons) from February 2005 to March 2013.

The most common analysis that looks at price relationships of markets primarily involves estimating the degree of market integration.[1] Less integrated markets—often indicated by significant and prolonged deviations from a stable, long-run relationship between these markets—reflect some form of market inefficiency. The markets are inefficient due to the fact that opportunities for spatial arbitrage exist that can still be exploited by traders.[2] Earlier methods that consider the degree of market integration use standard correlation coefficients and simple ordinary least squares (OLS) regressions.

[1] Goodwin and Piggott (2001) define market integration as "the extent to which shocks are transmitted among spatially separate markets." Since we are analyzing how price shocks in one market are transmitted to other markets, we will use price integration and market integration interchangeably. The concept of cointegration can also be applied. When variables are in a long-run equilibrium, the series are said to be cointegrated.

[2] A closely related concept that is also focused on spatial arbitrage is the law of one price (LOP). See Ardeni (1989), Baffes (1991), and Fackler and Tastan (2008) for discussions.

However, this approach has had widespread criticism, particularly with respect to a lack of accounting for the use of nonstationary data (Goodwin, Grennes, and Wohlgenant, 1990; Goodwin, 1992; Werden and Froeb, 1993). Stationarity (i.e., the probability distributions of a data are stable over time) is a critical property that makes regression techniques on time series data meaningful. Failing to adjust for nonstationary data would lead to what Granger and Newbold (1974) refer to as a "spurious regression problem": significant regression results may indicate the existence of a relationship between two variables when, in reality, there is no sense in which the relationship should exist in the first place.

The next wave of studies addressed this weakness by employing cointegration-based tests and time series regressions, particularly vector error correction models (Ravallion, 1986; Zanias, 1993; DeVany and Walls, 1993; Asche, Bremnes, and Wessells, 1999; Baulch, 1997a; Gulen, 1999; Kloveland, 2005). These new methods, based on the concept of cointegration as formally treated by Engle and Granger (1987), provided meaningful interpretations of results from regressions using nonstationary data. Applications of these techniques on agricultural commodities include analyses of the beef (Schroeder and Goodwin, 1990; Goodwin and Schroeder, 1991; Schroeder, 1997; Pendell and Schroeder, 2006) and pork industries (Faminow and Benson, 1990; Benson et al., 1994; Chen and Lee, 2008; Franken, Parcell, and Tonso, 2011). Analyses of market integration and spatial price asymmetry of oilseed and field crops, however, has not had much attention largely because oilseed and field crops tend to be harder to examine. Livestock markets are regional, and many local markets exist. On one hand, soybean oil, in particular, has fewer markets and these markets tend to have price movements typical of a national market. Furthermore, because of mandatory livestock price reporting, data on livestock volume movements tied to prices are available. As such, analyses of livestock price transmission that is typically motivated by market thinness, or concerns, is possible. On the other hand, co-reports of volume and price data are not readily available for oilseed crops, field crops, and their derived products. Instead, private data aggregators compile these data and sell them.

Goodwin (1992); Brester and Goodwin (1993); Kuiper, Lutz, and van Tilburg (1999); Gonzalez-Rivera and Helfand (2001); Goodwin and Piggott (2001); Thompson, Sul, and Bohl (2002); Franken et al. (2005); and Zhang, Escalante, and Wetzstein (2009) are among the few in the literature who have analyzed price relationships of oilseed crops, field crops, and their derived products. Studies that look at price relationships involving vegetable oils in particular include Duncker (1977); Labys (1977); Griffth and Meilke (1979); In and Inder (1997); Owen, Chowdhury, and Garrido (1997); Yu, Bessler, and Fuller (2006), and Peri and Baldi (2010). However, most of these studies focus on crosscountry analysis, and few, if any, focus on only one type of vegetable oil. Fewer still look at soybean oil, particularly for the United States.

In this paper, we aim to fill this literature gap by analyzing the relationship of soybean oil prices among four regional U.S. markets: Central Illinois, U.S. Gulf, West Coast, and East Coast. In particular, following the approach adopted by Franken et al. (2005; 2011), we examine whether the two types of U.S.-produced soybean oil—crude and RBD—exhibit long-run price relationships across these four markets and whether spatial pricing patterns have changed over time. Furthermore, we evaluate how the sudden biodiesel production increase during the mid-2000s might have caused changes in spatial price relationships among the four soybean oil markets.

We specifically emphasize the effects of biodiesel production because U.S.-produced biodiesel primarily uses soybean oil as a feedstock. Increasing biodiesel production and the complementary sudden soybean oil demand growth may create new spatial price relationships among different markets of the same vegetable oil commodity.[3] It may have generated new markets that have improved reliable price information availability across markets. Alternatively, it may have generated some "middle points" between markets, which can weaken the long-run relationship between spatially separated markets.
In this study, we hypothesize the following: (i) given the transferability and commodity nature of soybean oil, spatial soybean oil markets have remained price-integrated over time and (ii) the sudden biodiesel production increase during the mid-2000s and complementary soybean oil demand growth, which is caused by soybean oil being a major biodiesel feedstock, served as an exogenous shock that has affected price relationships. Biodiesel's introduction and production growth have generated new markets for soybean oil use, significantly affected demand, and altered spatial pricing relationships among existing soybean oil markets for conventional uses.

The implications of the results of our study have practical uses. If we find that the markets are strongly integrated in particular, then even in the absence of price information in one market, one can estimate the prices in that market (or changes in the prices thereof) by simply using price information from one of the other three markets. Another application from the analysis of the effects of introducing biodiesel production on long-run price relationships is in soybean oil price research. A researcher who is interested in monitoring soybean oil prices, for instance, might benefit more from focusing on data from a market that is least affected by structural changes based on the results of this study.

[3] Zilberman et al. (2013) provides an excellent discussion on how the introduction of biofuel affects food-commodity prices in particular.

Model

One of the most widely used spatial competitive equilibrium models was provided by Takayama and Judge (1964). It is also considered to be a basis for analyzing spatial market integration (Awokuse and Bernard, 2007).[4] An application of the Takayama-Judge model states that if trade occurs between two regions or markets, changes in one market's price should lead to an identical price response in the other market. Statistical and econometric techniques can, therefore, analyze the degree of integration between markets of identical products differentiated only by location.

In implementing the Takayama-Judge model empirically, the methods demonstrated here follow from Franken, Parcell, and Tonso (2011). They used standard time-series procedures to examine price linkages and price responsiveness among spatially dispersed hog markets. We use similar procedures to test whether soybean oil prices are cointegrated (i.e., have a long-run relationship) and will not diverge in the long run.

We begin by checking if the data used in the analysis are stationary or not. The Augmented Dickey-Fuller (ADF) test on price variables checks for a unit root's presence (Dickey and Fuller, 1979). If a unit root exists (i.e., time series is not stationary) in one or both variables being analyzed, then estimating any relationship between these variables would be meaningless and would result in what they term "a spurious regression." However, if two nonstationary variables are integrated of the same degree, performing regression analysis on both variables could be potentially meaningful, and the two variables can be recognized as cointegrated. The Engle-Granger two-step method can determine this (Engle and Granger, 1987). The method starts by estimating the relationship between two nonstationary price series by ordinary least squares (OLS):

$$Y_t = \alpha_0 + \alpha_1 X_t + e_t \tag{1}$$

where Y_t and X_t are individual nonstationary price series, α_0 and α_1 are intercept and slope coefficients, and e_t is the error term. Using the ADF test on e_t checks for presence of a unit root. If a unit root does not exist (i.e., e_t is stationary), then the two price series, Y_t and X_t, are cointegrated, and a long-run equilibrium relationship can be estimated. Note that the nonstationary series has to be integrated of the same degree.

The Engle-Granger method, however, is applicable only for bivariate equations. For models involving more than two variables, cointegration tests commonly employ the Johansen (1988) method to investigate the number of cointegrating vectors (i.e., long-run relationships). Specifically, if there are n prices with r cointegrating vectors, then $n - r$

[4] See Faminow and Benson (1990) and Franken et al. (2005) for a theoretical discussion on the Takayama-Judge model.

stochastic trends exist. Equivalently, if all price series exhibit the same stochastic trend, there must be $n-1$ cointegrating vectors, i.e., all prices are pairwise cointegrated. If more than one common trend exists, however, then the price series is not fully integrated. Correspondingly, the null hypothesis for both tests is that there are no more than r cointegrating vectors. Two types of Johansen tests exist. One uses the trace statistic, and the other uses the maximum eigenvalue. The alternative hypotheses are different for the two test types. For the trace test statistic, the alternative is that there exists more than r cointegration vectors. For the maximum eigenvalue test statistic, the alternative is that there are exactly $r+1$ cointegration vectors. Lütkepohl et al. (2001) found that there are no major differences between the two tests. As such, for convenience, we only employ the maximum eigenvalue test.

To account for the possibility that biodiesel's introduction caused a structural change in long-run price relationships or test for potential regime shifts, a set of residual-based cointegration tests, developed by Gregory and Hansen (1996), were estimated as well using OLS as follows:

$$(2) \qquad Y_t = \alpha_0 + \alpha_1 DUMMY_t + \alpha_2 X_t + \alpha_3 X_t DUMMY_t + e_t$$

where Y_t and X_t are defined as above; $DUMMY_t$ is a binary dummy defined as 1 following a significant increase in U.S. biodiesel production (May 2007) and 0 prior to this; α_0 and α_2 are the intercept and slope coefficients prior to the biodiesel production increase, respectively; and, α_1 and α_3 represent the changes in the intercept and the slope coefficients after the significant increase. As in (1), an ADF test for stationary of e_t from (2) is used to test for cointegration. However, standard ADF critical values are not appropriate for (2), and the appropriate critical values are reported in Gregory and Hansen (1996).

Estimating (1), which analyzes the whole sample period and (2) which takes into account any structural change stimulated by the sudden U.S. biodiesel production increase, enables testing of several hypotheses. First, if both specifications indicate that all prices are consistently cointegrated, then the surge in U.S. biodiesel production did not notably affect long-run equilibrium relationships among the markets. Second, coefficient estimates allow comparison of market price integration before and after the significant U.S. biodiesel production increase. For instance, if α_3 in (2) is statistically different from zero, then price relationships changed with the sudden increase in U.S. biodiesel production. If it is not statistically different from zero, then price relationships did not change. Furthermore, comparing estimates of α_2 with $(\alpha_2 + \alpha_3)$ in (2) would reveal whether prices move more or less on a one-for-one basis (i.e., perfectly integrated) after the sudden U.S. biodiesel production increase relative to before the biodiesel production growth.

Because we consider multiple price locations in our analysis, (1) and (2) are estimated as a special case of a vector autoregressive (VAR) specification:

$$\Delta Y_t = \alpha_0 + \alpha_1 DUMMY_t + \alpha_2 X_t + \alpha_3 X_t DUMMY_t + \sum_{k=1}^{K} \beta_{11}(k)\Delta Y_{t-k} + \sum_{k=1}^{K} \beta_{12}(k)\Delta X_{t-k} + \Omega_t \quad (3)$$

where t refers to time ($t = 1, 2, \ldots, T$) which, in our analysis, refers to months; K is the lag length; and Ω is an $n \times 1$ vector of normally distributed random errors. The specification of (3) allows for efficient standard errors and unbiased coefficients that will be used in running hypothesis tests of α_2 and ($\alpha_2 + \alpha_3$), while accounting for simultaneity between price locations. In particular, we want to test if both are not statistically different from unity, which would lend support to *full integration* (i.e., a one-for-one relationship) among these markets, both before and after the sudden U.S. biodiesel production increase.

As a final test for market integration, price relationships among market locations are further analyzed by investigating whether the speed of price responsiveness among locations differs before and after the sudden U.S. biodiesel production increase. To do this, an error correction VAR, or vector error correction (VEC) model, that incorporates the binary $DUMMY_t$ variable is estimated:

$$\Delta Y_t = \beta_0 + \beta_1 \hat{e}_{t-1} + \beta_2(\hat{e}_{t-1} \times DUMMY_t) + \sum_{k=1}^{K} \beta_{11}(k)\Delta Y_{t-k} + \sum_{k=1}^{K} \beta_{12}(k)\Delta X_{t-k} + \lambda_t \quad (4)$$

where variables and subscripts are as defined in (3), and λ is a $n \times 1$ vector of normally distributed random errors. If two markets are highly integrated, they would quickly return to long-run equilibrium after each has been pushed to disequilibrium following price shocks (Enders, 1995). In (4), β_1 measures the speed of adjustment or the one period lagged errors' effect on a relative price change for the entire sample period, and β_2 measures the change in the speed-of-adjustment's magnitude for a relative price change only during the time period after the sudden U.S. biodiesel production increase. The lagged error terms specified in (4) are obtained from the OLS estimation of (1). The next two terms are lagged price change variables following the standard VEC model. A speed-of-adjustment coefficient (β_1) close to one in absolute value indicates a quick adjustment to respond to equilibrium deviations, whereas a value near zero indicates a slow adjustment. If the sudden increase in U.S. biodiesel production improves reliable price information availability across markets, then (an adjusted or aggregate) speed-of-adjustment ($\beta_1 + \beta_2$) nearer to one in absolute value relative to β_1 should be expected. If, however, U.S. biodiesel's introduction has weakened the long-run relationship between spatially separated markets, then the value should be closer to zero.

Data

Data used in our analysis are monthly average soybean oil prices from four U.S. regional markets: Central Illinois, U.S. Gulf, West Coast, and East Coast. Data from February 2005 to March 2013 were included. Because traders aren't expected to react significantly to price shocks from another market within a day, using monthly data is a more reasonable frequency. Crude soybean oil and RBD soybean oil price data were obtained from The Jacobsen (Jacobsen, 2013). These data were divided into two time periods to account for a possible structural shift in the relationship among soybean oil prices due to the sudden U.S. biodiesel production increase during the mid-2000s. We assume a pre-biodiesel surge period from February 2005 to May 2007 and a post-biodiesel surge period from May 2007 to March 2013. Table 1 reports summary statistics of the data. It is apparent that, after the U.S. biodiesel production surge, the average and the standard deviation of soybean oil prices have increased.

Prior to the market integration analysis, the appropriate lag structure for the ADF tests and all subsequent models was determined by minimizing the Akaike Information Criteria (AIC). Table 2 shows that two lags exhibit the lowest AIC values for both soybean oil types. Our econometric model was, therefore, set to two lags.

Table 1. Summary Statistics

Soybean Oil Price (Crude, Cents per Pound)				
Market	Mean	Std. Dev.	Min.	Max.
Entire Period (February 2005 to March 2013)				
Central Illinois	39.57	12.23	21.32	63.54
U.S. Gulf	41.29	13.13	21.45	65.10
West Coast	40.65	12.44	22.17	66.79
East Coast	40.57	12.55	21.4	66.79
Pre-Biodiesel Production Surge (February 2005 to April 2007)				
Central Illinois	25.15	2.76	21.32	31.19
U.S. Gulf	25.18	2.86	21.45	31.34
West Coast	26.40	3.08	22.17	34.13
East Coast	25.91	3.15	21.40	33.51
Post-Biodiesel Production Surge (May 2007 to March 2013)				
Central Illinois	45.06	9.67	28.52	63.54
U.S. Gulf	47.42	9.89	30.32	65.10
West Coast	46.07	10.13	29.05	66.79
East Coast	46.14	10.02	29.05	66.79
Soybean Oil Price (Refined-Bleached-Deodorized, Cents per Pound)				
Entire Period (February 2005 to March 2013)				
Central Illinois	43.75	12.96	25.58	69.04
U.S. Gulf	45.58	13.88	24.86	70.36
West Coast	46.50	13.36	26.93	71.29
East Coast	45.92	13.16	26.62	70.79
Pre-Biodiesel Production Surge (February 2005 to April 2007)				
Central Illinois	29.21	2.36	25.58	34.12
U.S. Gulf	29.08	2.85	24.86	36.34
West Coast	30.41	2.57	26.93	36.27
East Coast	30.15	2.53	26.62	35.83
Post-Biodiesel Production Surge (May 2007 to March 2013)				
Central Illinois	49.27	10.88	31.52	69.04
U.S. Gulf	51.85	10.92	33.69	70.36
West Coast	52.61	10.36	34.86	71.29
East Coast	51.92	10.25	33.86	70.79

Note: Data covers 98 monthly average observations, from February 2005 to March 2013.

Table 2. Selection-Order Using Akaike Information Criterion

	Soybean Oil	
Lags	Crude	Refined, Bleached and Deodorized
0	15.37	14.83
1	8.14	9.06
2	7.87 *	9.03 *
3	8.04	9.13
4	8.08	9.21
5	8.06	9.23
6	8.14	9.08

Notes: * *Indicates optimal number of lag. There are 92 monthly observations used, from August 2005 to March 2013.*

Results

Determining market integration starts with the ADF tests on the price series. Table 3 shows that, based on the ADF tests, the data don't provide enough evidence to reject the null hypothesis of a unit root's existence. The price series are, therefore, nonstationary. However, long-run relationships among the price series can still be estimated as long as each one is integrated of the same degree. Table 4 presents the results of the ADF tests on the first difference of each series. The p-values show that the null hypothesis can be rejected and that the first-differenced series are each found to be stationary. The price data are, therefore, integrated of the same degree, order 1.

Table 3. Augmented Dickey-Fuller Test on Soybean Oil Data

Market	Test Statistic (Zt)	P-value
Crude (Central Illinois)	-1.89	0.34
Crude (U.S. Gulf)	-1.96	0.30
Crude (West Coast)	-1.90	0.33
Crude (East Coast)	-1.92	0.32
RBD (Central Illinois)	-1.89	0.34
RBD (U.S. Gulf)	-1.83	0.37
RBD (West Coast)	-1.82	0.37
RBD (East Coast)	-1.84	0.36

Notes: The null hypothesis is that a unit root exists, i.e. time series is not stationary. Test is performed with 2 lags.

We next determine whether the price series is cointegrated so that we can eventually analyze the nature of long-run relationships among the series. Because more than two price series are analyzed, the Johansen unrestricted cointegration rank statistics were used to test for cointegration during the entire period (Enders, 1995). Table 5 shares the crude soybean oil results, and Table 6 shares RBD soybean oil results. Trace statistics computed from characteristic roots (i.e., eigenvalues) reject the null hypothesis of no cointegrating vector for both soybean oil types. Hence, each market pair is deemed cointegrated, meaning that long-run price relationships do exist among these four markets and for both soybean oil types.

Table 4. Augmented Dickey-Fuller Test on First-Differenced Soybean Oil Data

Market	Test Statistic (Zt)	P-value
Crude (Central Illinois)	-4.36	0
Crude (U.S. Gulf)	-4.07	0
Crude (West Coast)	-4.08	0
Crude (East Coast)	-4.13	0
RBD (Central Illinois)	-4.20	0
RBD (U.S. Gulf)	-4.03	0
RBD (West Coast)	-4.27	0
RBD (East Coast)	-4.28	0

Notes: The null hypothesis is that a unit root exists, i.e. time series is not stationary. Test is performed with 2 lags.

Table 5. Johansen Unrestricted Cointegration Rank Test Statistics (Crude Soybean Oil)

Market Pairs	Eigenvalue	Maximum Eigenvalue Statistic
Central Illinois/U.S. Gulf	0.16 **	4.26
Central Illinois/West Coast	0.04 *	2.33
Central Illinois/East Coast	0.04 *	2.06
U.S. Gulf/West Coast	0.07 **	4.30
U.S. Gulf/East Coast	0.07 **	4.20
West Coast/East Coast	0.06 *	2.76

*Notes: * and ** indicate 0.05 level (5%) and 0.01 level (1%) of signficance, respectively. Lag length is set to 2. Maximum eigenvalue statistic critical values are 6.65 and 3.76 for the 1% and 5% levels of significance, respectively. The tests use 96 monthly observations.*

Table 6. Johansen Unrestricted Cointegration Rank Test Statistics (Refined, Bleached and Deodorized Soybean Oil)

Market Pairs	Eigenvalue	Maximum Trace Statistic
Central Illinois/U.S. Gulf	0.12 *	2.81
Central Illinois/West Coast	0.06 *	3.09
Central Illinois/East Coast	0.06 *	3.17
U.S. Gulf/West Coast	0.13 *	2.66
U.S. Gulf/East Coast	0.14 *	2.66
West Coast/East Coast	0.13 *	2.86

*Notes: * indicates 0.01 level (1%) of significance. Lag length is set to 2. Maximum eigenvalue statistic critical value is 6.65. The tests use 96 monthly observations.*

Using the cointegration rank test allowed us to analyze how the U.S. biodiesel production increase during the mid-2000s may have affected long-run relationships in soybean oil prices reported by different markets. The cointegration rank test results for the period before and after the sudden increase in biodiesel production are reported in Table 7 for crude oil and Table 8 for RBD oil. Three observations should be noted. First, the price series exhibits cointegrating relationships prior to the sudden U.S. biodiesel production increase. Second, the resulting eigenvalues still show that price series is cointegrated after the U.S. biodiesel production increase, except for one market pair. The cointegrating relationship between Central Illinois and the U.S. Gulf market prices disappear. Third, the significance level in rejecting the null hypothesis of no cointegration has weakened after the sudden U.S. biodiesel production increase for several market pairings. One implication from this particular observation, as well as from the disappearance of the cointegrating relationship in prices between the Central Illinois and U.S. Gulf markets, is that the sudden U.S. biodiesel production increase may have had weakened soybean oil spatial price relationships.

Table 7. Johansen Unrestricted Cointegration Rank Test Statistics (Crude Soybean Oil), Before and After Production Surge in Biodiesel

	Before Biodiesel		After Biodiesel	
Market Pairs	Eigenvalue	Maximum Eigenvalue Statistic	Eigenvalue	Maximum Eigenvalue Statistic
Central Illinois/U.S. Gulf	0.35 **	0.14	0.21	7.18
Central Illinois/West Coast	0.22 **	1.73	0.11 **	2.32
Central Illinois/East Coast	0.28 **	0.26	0.10 **	2.13
U.S. Gulf/West Coast	0.26 **	0.13	0.19 *	4.22
U.S. Gulf/East Coast	0.25 **	0.38	0.17 *	4.82
West Coast/East Coast	0.41 **	0.71	0.20 **	2.20

*Notes: * and ** indicate 0.05 level (5%) and 0.01 level (1%) of significance, respectively. Lag length is set to 2. Maximum eigenvalue statistic critical values are 6.65 and 3.76 for the 1% and 5% level of significance, respectively. Number of samples is 25 prior to production surge in biodiesel (May 2007), and 71 after the surge.*

Table 8. Johansen Unrestricted Cointegration Rank Test Statistics (Refined, Bleached and Deodorized Soybean Oil), Before and After Production Surge in Biodiesel

	Before Biodiesel		After Biodiesel	
Market Pairs	Eigenvalue	Maximum Eigenvalue Statistic	Eigenvalue	Maximum Eigenvalue Statistic
Central Illinois/U.S. Gulf	0.19 **	3.18	0.12 *	4.45
Central Illinois/West Coast	0.30 **	0.59	0.15 **	3.52
Central Illinois/East Coast	0.30 **	0.15	0.15 **	3.51
U.S. Gulf/West Coast	0.15 **	1.30	0.17 **	3.64
U.S. Gulf/East Coast	0.17 **	1.67	0.16 *	3.77
West Coast/East Coast	0.37 **	0.19	0.13 *	5.45

*Notes: * and ** indicate 0.05 level (5%) and 0.01 level (1%) of significance, respectively. Lag length is set to 2. Maximum eigenvalue statistic critical values are 6.65 and 3.76 for the 1% and 5% level of significance, respectively. Number of samples is 25 prior to production surge in biodiesel (May 2007), and 71 after the surge.*

Table 9. VAR Parameter Estimates from Regime Shift Model (Crude Soybean Oil)

Dependent Market/ Independent Market	Constant (α_0)	Post-Biodiesel Dummy ($\alpha 1$)	State (α_2)	Post-Biodiesel Regime (α_3)	H_0: $\alpha_2 = 1$ (p-value)	H_0: $\alpha_2 + \alpha_3 = 1$ (p-value)
Central Illinois/U.S. Gulf	0.07	− 1.06	0.97 *	0.01	0.47	0.37
	(0.87)	(0.92)	(0.04)	(0.03)		
U.S. Gulf/Central Illinois	− 0.07	1.08	1.02 *	− 0.01	0.51	0.42
	(0.92)	(0.96)	(0.04)	(0.03)		
Central Illinois/West Coast	0.84	− 0.03	0.97 *	0.02	0.38	0.58
	(0.86)	(0.90)	(0.03)	(0.03)		
West Coast/Central Illinois	− 0.87	0.06	1.03 *	− 0.02	0.41	0.62
	(0.92)	(0.96)	(0.04)	(0.04)		
Central Illinois/East Coast	0.36	0.23	0.98 *	0.01	0.44	0.28
	(0.80)	(0.84)	(0.03)	(0.03)		
East Coast/Central Illinois	− 0.38	− 0.23	1.02 *	− 0.01	0.50	0.31
	(0.89)	(0.94)	(0.04)	(0.04)		
U.S. Gulf/West Coast	1.08	0.72	0.98 *	0.01	0.66	0.79
	(1.18)	(1.23)	(0.05)	(0.04)		
West Coast/U.S. Gulf	− 1.08	− 0.72	1.02 *	− 0.01	0.72	0.89
	(1.21)	(1.27)	(0.05)	(0.05)		
U.S. Gulf/East Coast	0.67	0.90	0.98 *	< 0.00	0.71	0.55
	(1.10)	(1.16)	(0.04)	(0.04)		
East Coast/U.S. Gulf	− 0.67	− 0.91	1.01 *	< − 0.00	0.79	0.64
	(1.18)	(0.1.24)	(0.05)	(0.05)		
West Coast/East Coast	− 0.11	− 0.12	1.00 *	< − 0.00	0.89	0.41
	(0.16)	(0.17)	0.00	(0.01)		
East Coast/West Coast	0.12	0.12	1.00 *	< 0.00	0.89	0.41
	(0.16)	(0.17)	(0.01)	(0.01)		

*Notes: * indicates 0.01 level (1%) of significance. Standard errors in parenthesis. Lag length is set to 2. Number of samples is 92.*

Following Pendell and Schroeder (2006), the VAR model (3) was next estimated to test the strength of the price linkages between the four markets and verify that the prices do not diverge from one another in the long run. A dummy variable was used to represent the timing of the sudden U.S. biodiesel production increase (= 1 after May 2007, = 0 otherwise). The results lend support to full integration among these markets, either before or after the sudden U.S. biodiesel production increase. See Table 9 for crude soybean oil and Table 10 for RBD soybean oil. Specifically, not enough evidence is available to reject the null hypothesis that the price coefficient (α_2) equals one. This result indicates full price integration prior to the U.S. biodiesel production increase. The same is true for the null hypothesis that the sum of the price coefficient and the dummy interaction term ($\alpha_2 + \alpha_3$) equals one, indicating full integration after the sudden increase. This finding is consistent across alternative market pairings. In summary, the VAR model's results indicate that soybean oil prices in the four markets were fully integrated both before the U.S. biodiesel production increase and after it.

Speed-of-adjustment coefficients from the VEC model (4) are reported in Table 11 for crude soybean oil and Table 12 for RBD soybean oil. Recall that if the sudden U.S. biodiesel production increase improves reliable price information availability, then the adjusted or aggregate speed-of-adjustment measure ($\beta_1 + \beta_2$) should be nearer to one in absolute value than the simple, unadjusted measure (β_1). We find, however, that this relationship doesn't occur consistently in all cases for both soybean oil types. Take crude soybean oil in the Central Illinois market as an example. Its price relationship improved after the U.S. biodiesel production surge only with the U.S. Gulf market. In response to a one unit deviation from equilibrium in period $t - 1$, the Central Illinois price falls by 0.5271 units, and the U.S. Gulf price rises by 0.3535 units. Both changes are larger than the degree of adjustment before the marked biodiesel production increase (–0.2448 and –0.0180, respectively). For Central Illinois' pairings with the West Coast and East Coast markets, the price adjustments were faster prior to the U.S. biodiesel production surge. Nevertheless, despite the mixed results, an important observation common to all market pairings is that the differences in price adjustments before and after the surge in U.S. biodiesel production are clearly distinct and, in most cases, large. This indicates that the U.S. biodiesel production surge has caused significant changes in spatial price relationships that may have eventually led to either substantial increases or decreases in speed-of-price adjustments toward equilibrium.

The sudden increase in U.S. biodiesel production's effect is also evident when looking at Granger causality tests for RBD soybean oil.[5] The Granger causality test is another useful tool in analyzing the price relationships among the four markets because it

[5] Granger causality tests for crude soybean oil yielded no significant differences in the results before and after the U.S. biodiesel production surge in the mid-2000s.

examines whether the future prices from one market can consistently be better predicted using historical prices from another market (Granger, 1969). Table 13 presents Granger causality test results corresponding to the VEC framework to determine the extent to which lagged prices for one RBD soybean oil market influence prices in another market. Test statistics for the null hypothesis of no causality are presented for portions of the sample before and after the U.S. biodiesel production increase, as well as the entire sample period. All but two market pairings show two-way Granger causality in prices before the U.S. biodiesel production increase. The exceptions are prices in the Central Illinois market, which Granger cause prices in the West Coast and East Coast markets, but not the other way around. All Granger causalities cease to exist, however, after the production increase. These findings convey two things. First, the contrasting Granger causality results between the two time periods suggest that the increase in U.S. biodiesel production may have caused a structural change in the price relationships among the four markets. Second, the disappearance of any Granger causalities after the production increase indicates that such structural change weakened the price relationships. Prices from any one market no longer have predictive power in forecasting prices from any other market after the sudden increase in U.S. biodiesel production.

Table 10. VAR Parameter Estimates from Regime Shift Model (Refined-Bleached- Deodorized Soybean Oil)

Dependent Market/ Independent Market	Constant (α_0)	Post-Biodiesel Dummy ($\alpha 1$)	State (α_2)	Post-Biodiesel Regime (α_3)	H_0: $\alpha_2 = 1$ (p-value)	H_0: $\alpha_2 + \alpha_3 = 1$ (p-value)
Central Illinois/U.S. Gulf	1.09	– 1.65 *	0.97 **	0.04	0.41	0.55
	(0.90)	(0.96)	(0.03)	(0.03)		
U.S. Gulf/Central Illinois	– 1.21	1.75	1.03 **	– 0.04	0.43	0.48
	(1.04)	(1.07)	(0.04)	(0.03)		
Central Illinois/West Coast	0.41	– 2.14	0.98 **	0.04	0.65	0.43
	(1.53)	(1.62)	(0.05)	(0.05)		
West Coast/Central Illinois	– 0.44	2.15	1.02 **	– 0.04	0.66	0.37
	(1.57)	(1.64)	(0.05)	(0.05)		
Central Illinois/East Coast	0.32	– 1.95	0.99 **	0.04	0.86	0.18
	(1.51)	(1.59)	(0.05)	(0.05)		
East Coast/Central Illinois	– 0.34	1.93	1.01 **	– 0.04	0.86	0.15
	(1.52)	(1.59)	(0.05)	(0.05)		
U.S. Gulf/West Coast	– 1.51	0.48	1.03 **	– 0.03	0.53	0.78
	(1.53)	(1.57)	(0.05)	(0.05)		
West Coast/U.S. Gulf	1.43	– 0.38	0.97 **	0.02	0.48	0.67
	(1.36)	(1.37)	(0.05)	(0.05)		
U.S. Gulf/East Coast	– 1.59	0.66	1.04 **	– 0.03	0.42	0.61
	(1.53)	(1.57)	(0.05)	(0.05)		
East Coast/U.S. Gulf	1.48	– 0.54	0.96 **	0.03	0.36	0.50
	(1.33)	(1.35)	(0.05)	(0.04)		
West Coast/East Coast	– 0.17	0.20	1.02 **	< – 0.00	0.26	0.06
	(0.42)	(0.44)	(0.01)	(0.01)		
East Coast/West Coast	0.17	– 0.20	0.98 **	< 0.00	0.25	0.05
	(0.42)	(0.43)	(0.01)	(0.01)		

Notes: * *and* ** *indicate 0.10 level (10%) and 0.01 level (1%) of significance, respectively. Standard errors in parenthesis. Lag length is set to 2. Number of samples is 92.*

Table 11. Speed-of-Adjustment Coefficients from Vector Error Correction Model (Crude Soybean Oil)

Dependent Market/ Independent Market	Speed-of-Adjustement Coefficient (Entire Period) (β_1)	Size of Speed-of-Adjustment after Biodiesel (β_2)	Net Impact ($\beta_1 + \beta_2$)
Central Illinois/U.S. Gulf	− 0.24 (0.53)	− 0.28 (0.57)	− 0.53
U.S. Gulf/Central Illinois	− 0.02 (0.47)	0.37 (0.53)	0.35
Central Illinois/West Coast	0.09 (0.47)	− 0.01 (0.50)	0.08
West Coast/Central Illinois	− 0.57 (0.53)	0.45 (0.55)	− 0.12
Central Illinois/East Coast	− 0.09 (0.51)	0.16 (0.55)	0.07
East Coast/Central Illinois	− 0.34 (0.52)	0.22 (0.55)	− 0.12
U.S. Gulf/West Coast	0.08 (0.31)	0.03 (0.34)	0.11
West Coast/U.S. Gulf	− 0.43 (0.41)	0.27 (0.43)	− 0.15
U.S. Gulf/East Coast	0.03 (0.34)	0.10 (0.37)	0.13
East Coast/U.S. Gulf	− 0.40 (0.43)	0.20 (0.45)	− 0.19
West Coast/East Coast	− 0.16 (1.27)	− 0.16 (1.54)	− 0.32
East Coast/West Coast	0.03 (1.25)	0.23 (1.53)	0.26

Notes: Net impact may not add up due to rounding. Standard errors in parenthesis. Lag length is set to 2. Number of samples is 95.

Table 12. Speed-of-Adjustment Coefficients from Vector Error Correction Model (Refined, Bleached and Deodorized Soybean Oil)

Dependent Market/ Independent Market	Speed-of-Adjustement Coefficient (Entire Period) (β_1)	Size of Speed-of-Adjustment after Biodiesel (β_2)	Net Impact ($\beta_1 + \beta_2$)
Central Illinois/U.S. Gulf	− 0.27 (0.40)	− 0.16 (0.47)	− 0.43
U.S. Gulf/Central Illinois	0.07 (35.00)	0.22 (0.43)	0.29
Central Illinois/West Coast	− 0.26 (0.53)	0.11 (0.59)	− 0.15
West Coast/Central Illinois	0.16 (0.45)	− 0.16 (0.52)	< − 0.00
Central Illinois/East Coast	− 0.25 (0.58)	0.10 (0.64)	− 0.15
East Coast/Central Illinois	0.14 (0.49)	− 0.15 (0.57)	− 0.01
U.S. Gulf/West Coast	− 0.13 (0.57)	0.22 (0.61)	0.10
West Coast/U.S. Gulf	− 0.09 (0.57)	− 0.30 (0.61)	− 0.40
U.S. Gulf/East Coast	− 0.10 (0.54)	0.20 (0.58)	0.11
East Coast/U.S. Gulf	− 0.20 (0.55)	− 0.19 (0.59)	− 0.39
West Coast/East Coast	2.90 (3.14)	− 2.82 (3.26)	0.08
East Coast/West Coast	− 3.39 (3.22)	3.04 (3.33)	− 0.35

Notes: Net impact may not add up due to rounding. Standard errors in parenthesis. Lag length is set to 2. Number of samples is 95.

Table 13. Granger Causality for Soybean Oil Prices from Vector Error Correction Model (Refined, Bleached and Deodorized Soybean Oil)

Dependent Market/	χ^2 Test Statistic		
Independent Market	Pre-Biodiesel	Post-Biodiesel	Entire Period
Central Illinois/U.S. Gulf	14.79 **	2.53	1.07
Central Illinois/West Coast	2.93	2.12	1.41
Central Illinois/East Coast	3.72	1.62	1.15
U.S. Gulf/Central Illinois	11.07 **	1.77	0.97
U.S. Gulf/West Coast	12.83 **	1.65	0.87
U.S. Gulf/East Coast	9.04 *	1.50	0.77
West Coast/Central Illinois	6.60 *	4.10	4.48
West Coast/U.S. Gulf	9.83 **	4.37	3.66
West Coast/East Coast	6.81 *	0.95	0.46
East Coast/Central Illinois	8.13 *	3.31	4.03
East Coast/U.S. Gulf	7.57 *	4.43	3.58
East Coast/West Coast	6.36 *	1.22	0.71

*Notes: * and ** indicate 0.05 level (5%) and 0.01 level (1%) of significance, respectively. Lag length is set to 2.*

Conclusion

We have investigated the price relationships among four U.S. soybean oil markets and analyzed whether the sudden U.S. biodiesel production increase during the mid-2000s has had any detectable impact on price relationships among regional markets. Our results show that crude and RBD soybean oil prices from four regional markets—Central Illinois, U.S. Gulf, West Coast, and East Coast—are cointegrated, showing that prices share long-run relationships. VAR results also provide strong evidence that the markets are fully integrated. However, the Johansen cointegration tests, the speed-of-adjustment coefficients from the VEC model, and the Granger causality test results show that the sudden U.S. biodiesel production increase during the mid-2000s may have changed and, to some extent, may have weakened the spatial price relationships.

Thus, the four soybean oil markets being analyzed are found to have remained price-integrated over time. This finding implies that the markets are spatially efficient. Any slight divergence in prices leads to arbitrage that would make markets adjust quickly and prevent prices from deviating farther. The results, however, also suggest that the level of market efficiency may have decreased, only marginally, after the U.S. biodiesel production surge. One possible source is that the entry of the biodiesel industry and the complementary soybean oil demand growth caused by soybean oil being used as a major biodiesel feedstock have created new markets that siphon soybean oil towards production of biodiesel. The creation of these new markets significantly affected demand and may have caused lowering of the speed of the price adjustments among the four major soybean oil markets.

For future research, this paper can be extended by addressing three important issues. First, cointegration-based tests have been criticized in recent literature due to the fact that their procedures ignore transaction costs (Barrett, 1996; Goodwin and Piggott, 2001). Balke and Fomby (1997) discussed the tendency for two variables to move toward a long-run equilibrium that may not occur in every period and that there may be a certain threshold where cointegration is triggered.[6] Transaction costs or policy interventions may create a band within which the two variables are not cointegrated. More recent papers attempt to address this by employing the latest econometric methods used in spatial price analysis, such as threshold analyses and the endogenous switching models (Spiller and Wood, 1988; Sexton, Kling, and Carman, 1991; Baulch, 1997b; Balcombe, Bailey, and Brooks, 2007). In our analysis, however, the fact that the markets still generally exhibit full integration may reflect the absence of transaction costs. According to Franken et al. (2011), this circumstance can justify not using these new methods.

Second, the results of our analysis may relate to the efficient market hypothesis (EMH). Contrary to the implication of the results discussed, EMH posits that the existence of cointegration of prices across markets indicates market inefficiency. This is based on the idea that markets are (information-) efficient if the prices are determined independently (Fama, 1970). Being able to forecast price from one market using historical prices from another market violates efficiency. Little work has been done in the literature that uses the EMH framework to test the efficiency of U.S. agricultural commodity markets. Among those is Yang and Leatham (1998), which tests the market efficiency of the U.S. grain markets. While our paper looks at the price relationships of one agricultural commodity among different geographical markets (soybean oil), Yang and Leatham looks at the price relationships among four commodities (corn, oats, wheat, and soybeans). Reconciling the EMH with the results of our empirical analysis is an important issue that must be addressed at the theoretical level.

Third, other factors occurring in the industry may also affect the price relationships. However, we didn't account for such factors mostly because adequate data weren't available to consider these relationships. Two factors in particular are worth noting. First,

[6] Extensive research has already been done applying the threshold cointegration approach. See Heckscher (1916), Obstfeld and Taylor (1997), Goodwin and Schroeder (1991), Goodwin and Grennes (1998), Goodwin and Piggott (2001), Meyer (2004).

consumer demand for healthy oils decreased domestic edible soybean oil's market share by 16 percentage points during the previous decade. To extend shelf-life and achieve frying stability, soybean oil is typically hydrogenized, partially or wholly, and the process adds trans fats, which consumers and food companies have attempted to exclude from their diets or formulations, respectively. Second, global economic wealth expansion has increased during the past decade and boosted world soybean oilseed exports from 53.82 million metric tons in 2000-01 to 92.27 million metric tons in 2011-12 (USDA, 2013b). These two events may have affected spatial price relationships of soybean oil markets, but accounting for these factors is beyond the scope of our analysis.

References

Ardeni, P.G. (1989). "Does the Law of One Price Really Hold for Commodity Prices?" *American Journal of Agricultural Economics* 71(3): 661-669.

Asche, F., H. Bremnes, and C.R. Wessells. (1999). "Product Aggregation, Market Integration, and Relationships Between Prices: An Application to World Salmon Markets." *American Journal of Agricultural Economics* 81(3): 568-581.

Awokuse, T.O., and J.C. Bernard. (2007). "Spatial Price Dynamics in U.S. Regional Broiler Markets." *Journal of Agricultural and Applied Economics* 39(3): 447-456.

Balcombe, K., A. Bailey, and J. Brooks. (2007). "Threshold Effects in Price Transmission: The Case of Brazilian Wheat, Maize, and Soya Prices." *American Journal of Agricultural Economics* 89(2): 308-323.

Balke, N.S., and T.B. Fomby. (1997). "Threshold Cointegration." *International Economic Review* 38(3): 627-645.

Barrett, C.B. (1996). "Market Analysis Methods: Are Our Enriched Toolkits Well-Suited to Enlivened Markets?" *American Journal of Agricultural Economics* 78(3): 825-829.

Baulch, B. (1997a). "Testing for Food Market Integration Revisited." *The Journal of Development Studies* 33(4): 512-534.

Baulch, B. (1997b). "Transfer Costs, Spatial Arbitrage, and Testing for Food Market Integration." *American Journal of Agricultural Economics* 79(2): 477-487.

Baffes, J. (1991). "Some Further Evidence on the Law of One Price: The Law of One Price Still Holds." *American Journal of Agricultural Economics* 73(4):1264-1273.

Benson, B.L., M.D. Faminow, M.H. Marquis, and D.G. Sauer. (1994). "Intra-national Effects of a Countervailing Duty on the United States/Canadian Hog Market." *Review of Agricultural Economics* 16(2): 187-201.

Braun, J.V. (2008). "Responding to the World Food Crisis: Getting on the Right Track." in *Responding to the Global Food Crisis: Three Perspectives*. International Food Policy Research Institute. Washington, D.C.

Brester, G.W., and B.K. Goodwin. (1993). "Vertical and horizontal Price Linkages and Market Concentration of the U.S. Wheat Milling Industry." *Review of Agricultural Economics* 15(3): 507-519.

Chen, P.F., and C.C. Lee. (2008). "Nonlinear Adjustments in Deviations from the Law of One Price for Wholesale Hog Prices." *Agricultural Economics* 39(1): 123-134.

DeVany, A., and W.D. Walls. (1993). "Pipeline Access and Market Integration in the Natural Gas Industry: Evidence from Cointegration Tests." *The Energy Journal* 14(4): 1-20.

Dickey, D.A., and W.A. Fuller. (1979). "Distribution of the Estimators for Autoregressive Time Series With a Unit Root." *Journal of American Statistical Association* 74(366a): 427-431.

Duncker, J.W. (1977). "World Price Behaviour of Edible Oils and Fats." *Canadian Farm Economics* 12: 14-22.

Enders, W. (1995). *Applied Econometric Time Series*. New York, NY: John Wiley & Sons, Inc.

Engle, R.F., and C.W.J. Granger. (1987). "Co-integration and Error Correction: Representation, Estimation, and Testing." *Econometrica* 55(2): 251-276.

Fackler, P.L., and H. Tastan. (2008). "Estimating the Degree of Market Integration." *American Journal of Agricultural Economics* 90(1): 69-85.

Fama, E.F. (1970). "Efficient Capital Markets: A Review of Theory and Empirical Work." *Journal of Finance* 25(2): 383-417.

Faminow, M.D., and B.L. Benson. (1990). "Integration of Spatial Markets." *American Journal of Agricultural Economics* 72(1): 49-62.

Franken, J.R.V., J.L. Parcell, M.E. Sykuta, and C.L. Fulcher. (2005). "Market Integration: Case Studies of Structural Change." *Agricultural and Resource Economics Review* 34(2): 163-172.

Franken, J.R.V., J.L. Parcell, and G.T. Tonsor. (2011). "Impact of Mandatory Price Reporting on Hog Market Integration." *Journal of Agricultural and Applied Economics* 43(2): 229-241.

Gonzalez-Rivera, G., and S.M. Helfand. (2001). "The Extent, Pattern and Degree of Market Integration: A Multivariate Approach for the Brazilian Rice Market." *American Journal of Agricultural Economics* 83(3): 576-592.

Goodwin, B.K. (1992). "Mutivariate Cointegration Tests and the Law of One Price in International Wheat Markets." *Review of Agricultural Economics* 14(1): 117-124.

Goodwin, B.K., and T.J. Grennes. (1998). "Tsarist Russia and the World Wheat Market." *Explorations in Economic History* 35(4): 405-430.

Goodwin, B.K., and N.E. Piggott. (2001). "Spatial Market Integration in the Presence of Threshold Effects." *American Journal of Agricultural Economics* 83(2):302-317.

Goodwin, B.K., and T.C. Schroeder. (1991). "Cointegration Tests and Spatial Price Linkages in Regional Cattle Markets." *American Journal of Agricultural Economics* 73(2): 453-464.

Goodwin, B.K., T.J. Grennes, and M.K. Wohlgenant. (1990). "A Revised Test of the Law of One Price Using Rational Price Expectations." *American Journal of Agricultural Economics* 72(3): 682-693.

Granger, C.W.J. (1969). "Investigating Causal Relations by Econometric Models and Cross-Spectral Methods." *Econometrica* 37(3): 424-438.

Granger, C.W.J., and P. Newbold. (1974). "Spurious Regressions in Econometrics." *Journal of Econometrics* 2(2): 111-120.

Gregory, A.W., and B.E. Hansen. (1996). "Residual-Based Tests for Cointegration in Models with Regime Shifts." *Journal of Econometrics* 70(1): 99-126.

Griffith, G.R., and K.D. Meilke. (1979). "Relationships Among North American Fats and Oils Prices." *American Journal of Agricultural Economics* 61(2): 335-341.

Gulen, S.G. (1999). "Regionalization in the World Crude Oil Market: Further Evidence." *Energy Journal* 20(1): 125-139.

Heckscher, E.F. (1916). "Växelkursens Grundval vid Pappersmyntfot." *Ekonomisk Tidskrift* 18(10): 309-312.

In, F., and B. Inder. (1997). "Long-Run Relationships Between World Vegetable Oil Prices." *Australian Journal of Agricultural and Resource Economics* 41(4): 455-470.

Jacobsen. (2013). *Vegetable Oils*. The Jacobsen, Chicago, IL. Available at http://www.thejacobsen.com.

Johansen, S. (1988). "Statistical Analysis of Cointegrating Vectors." *Journal of Economic Dynamics and Control* 12(2,3): 213-254.

Kloveland, J.T. (2005). "Commodity Market Integration 1850-1913: Evidence from Britain and Germany." *European Review of Economic History* 9(2): 163-197.

Kuiper, W.E., C. Lutz, and A. van Tilburg. (1999). "Testing for the Law of One Price and Identifying Price-Leading Markets: An Application to Corn Markets in Benin." *Journal of Regional Science* 39(4): 713-738.

Labys, W.C. (1977). "Mutlicommodity Substitution Patterns in the International Fats and Oils Market." *European Review of Agricultural Economics* 4(1): 75-84.

Lütkepohl, H., P. Saikkonen, and C. Trenkler. (2001). "Maximum Eigenvalue Versus Trace Tests for the Cointegrating Rank of a VAR Process." *The Econometrics Journal* 4(2): 287-310.

Meyer, J. (2004). "Measuring Market Integration in the Presence of Transaction Costs—a Threshold Vector Error Correction Approach." *Agricultural Economics* 31(2,3): 327-334.

Mittal, A. (2009). "The 2008 Food Price Crisis: Rethinking Food Security Policies." G-24 Discussion Paper no. 56, United Nations Conference on Trade and Development.

Obstfeld, M., and A.M. Taylor. (1997). "Nonlinear Aspects of Goods-Market Arbitrage and Adjustment: Heckscher's Commodity Points Revisited." *Journal of the Japanese and International Economies* 11(4): 441-479.

Owen, A.D., K. Chowdhury, and J.R.R. Garrido. (1997). "Price Interrelationship in the Vegetable and Tropical Oils Market." *Applied Economics* 29(1): 119-124.

Pendell, D.L., and T.C. Schroeder. (2006). "Impact of Mandatory price Reporting on Fed Cattle Market Integration." *Journal of Agricultural and Resource Economics* 31(3): 568-579.

Peri, M., and L. Baldi. (2010). "Vegetable Oil Market and Biofuel Policy: An Asymmetric Cointegration Approach." *Energy Economics* 32(3): 687-693.

Ravallion, M. (1986). "Testing Market Integration." *American Journal of Agricultural Economics* 68(1): 292-307.

Schroeder, T.C. (1997). "Fed Cattle Spatial Transactions Price Relationships." *Journal of Agricultural and Applied Economics* 29(2): 347-362.

Schroeder, T.C., and B.K. Goodwin. (1990). "Regional Fed Cattle Price Dynamics." *Western Journal of Agricultural Economics* 15(1): 111-122.

Sexton, R.J., C.L. Kling, and H.F. Carman. (1991). "Market Integration, Efficiency of Arbitrage, and Imperfect Competition: Methodology and Application to U.S. Celery." *American Journal of Agricultural Economics* 73(3): 568-580.

Spiller, P.T., and R.O. Wood. (1988). "The Estimation of Transaction Costs in Arbitrage Models." *Journal of Econometrics* 39(3): 309-326.

Steinberg, S. (2008). "Financial Speculators Reap Profits from Global Hunger" *Global Research*, Centre for Research on Globalization, April 24. Available at http://www.globalresearch.ca/financial-speculators-reap-profits-from-global-hunger/8794 (accessed December 29, 2013).

Takayama, T., and G.G. Judge. (1964). "Spatial equilibrium and Quadratic Programming." *Journal of Farm Economics* 46(1): 67-93.

Thompson, S.R., D.Y. Sul, and M.T. Bohl. (2002). "Spatial Market Efficiency and Policy Regime Change: Seemingly Unrelated Error Correction Model Estimation." *American Journal of Agricultural Economics* 84(4): 1041-1053.

U.S. Department of Agriculture. (2013a). *Economic Research Service Oil Crops Yearbook*. U.S. Department of Agriculture, Washington, D.C. Available at http://www.ers.usda.gov/data-products/oil-crops-yearbook.aspx (accessed August 25, 2013).

U.S. Department of Agriculture. (2013b). *Economics, Statistics, and Market Information System*. U.S. Department of Agriculture, Washington, D.C. Available at http://usda.mannlib.cornell.edu/MannUsda/homepage.do.

U.S. Energy Information Administration. (2013a). *Annual Energy Review*. U.S. Energy Information Administration, Washington, D.C. Available at http://www.eia.gov/totalenergy/data/annual/index.cfm.

U.S. Energy Information Administration. (2013b). *Monthly Biodiesel Production Report*. U.S. Energy Information Administration, Washington, D.C. Available at http://www.eia.gov/biofuels/biodiesel/production.

Werden, G.J., and L.M. Froeb. (1993). "Correlation, Causality, and All that Jazz: The Inherent Shortcomings of Price Tests for Antitrust Market Delineation." *Review of Industrial Organization* 8(3): 329-353.

Yang, J., and D.J. Leatham. (1998). "Market Efficiency of US Grain Markets: Application of Cointegration Tests." *Agribusiness* 14(2): 107-112.

Yu, T.H., D. A. Bessler, and S. Fuller. (2006). "Cointegration and Causality Analysis of World Vegetable Oil and Crude Oil Prices." Selected paper presented at the 2006 annual meeting of the American Agricultural Economics Association, Long Beach, CA.

Zanias, G.P. (1993). "Testing for Integration in European Community Agricultural Product Markets." *Journal of Agricultural Economics* 44(3): 418-427.

Zhang, Z., L. Lohr, C. Escalante, and M. Wetzstein. (2009). "Ethanol, Corn, and Soybean price Relations in a Volatile Vehicle-Fuels Market." *Energies* 2(2): 320-339.

Zilberman, D., G. Hochman, D. Rajagopal, S. Sexton, and G. Timilsina. (2013). "The Impact of Biofuels on Commodity Food Prices: Assessment of Findings." *American Journal of Agricultural Economics* 95(2): 275-281.