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Soybean Oil Spatial Price Dynamics

We analyze the price relationship of 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 the level of market efficiency may have decreased to some extent after the U.S. biodiesel production surge in the mid-2000s.

Keywords: spatial price analysis, soybean oil, market integration, vector autoregression, vector error correction, biodiesel

Introduction

Vegetable oils are used for food, cooking, biofuel production, and industrial uses. These oils are derived from processing oilseeds chemically or using a press. The resulting oil is typically 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 U.S. 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 percent, in 2011 (USDA, 2013).

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

Figures 1 display historical price trends for refined-bleached-deodorized (RBD) soybean oils. This figure indicates 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 percent, rice prices increased by 170 percent 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 million gallons in 2003 , output jumped to an average of 40 million gallons in 2007 (Figure 2). These two events were also closely related. The International Food Policy Research Institute (IFPRI) calculated that 30 percent of the increase in average grain prices between 2000 and 2007 was accounted for by biofuel production (Braun, 2008).

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 is especially notable after 2007. Although the spreads may appear minimal, even small deviations can signal considerable spatial price arbitrage opportunities. Central Illinois RBD soybean oil prices, for example, 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. 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.

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 long-run relationships between markets, reflect some form of market inefficiency, suggesting opportunities for spatial arbitrage.[2] Earlier methods that consider the degree of market integration use standard correlation coefficients and simple ordinary least squares (OLS) regressions. However, this approach has had widespread criticism, particularly with respect to lack of accounting for the use of non-stationary data (Goodwin et al., 1990; Goodwin, 1992; Werden and Froeb, 1993).

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 et al., 1999; Baulch, 1997a; Gulen, 1999; Kloveland, 2005). Applications of these techniques on agricultural commodities include analyses of beef industries (Schroeder and Goodwin, 1990; Goodwin and Schroeder, 1991; Schroeder, 1997; Pendell and Schroeder, 2006) and pork industries (Faminov and Benson, 1990; Benson et al., 1994; Chen and Lee, 2008; Franken et al., 2011). Analyses of market integration and spatial price asymmetry of oilseed and field crops, however, relatively has not had much attention because oilseed and field crops tend to be harder to examine. Livestock markets are regional, and many local markets exist. For soybean oil in particular, on the other hand, fewer markets exist 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 are typically motivated by market thinness, or concerns, are 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 et al. (1999), Gonzalez-Rivera and Helfand (2001), Goodwin and Piggott (2001), Thompson et al. (2002), Franken et al. (2005), and Zhang et al. (2009) are among the few in the literature that 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

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[1] Goodwin and Piggott (2001) defines market integration as "the extent to which shocks are transmitted among spatially separate markets." Since we are analyzing how price shocks in one market is 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.

(1979), In and Inder (1997), Owen et al. (1997), Yu et al. (2006), and Peri and Baldi (2010). However, most of these studies focus on cross-country analysis, and few, if any, focus on only one type of vegetable oil. Fewer still look at soybean oil, particularly for the U.S.

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 U.S.-produced RBD soybean oil 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.

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 base 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 et al. (2011). Franken et al. 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.

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[3] Zilberman et al. (2013) provides an excellent discussion on how the introduction of biofuel affects food-commodity prices in particular.
[4] See Faminow and Benson (1990) and Franken et al. (2005) for a theoretical discussion on the Takayama-Judge model.

To begin, 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 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 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 $n - r$ 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 are 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, 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.

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 using OLS as follows:

$$Y_t = \alpha_0 + \alpha_1 DUMMY_t + \alpha_2 X_t + \alpha_3 X_t DUMMY_t + e_t \tag{2}$$

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:

$$\begin{aligned} \Delta Y_t = {} & \alpha_0 + \alpha_1 DUMMY_t + \alpha_2 X_t + \alpha_3 X_t DUMMY_t \\ & + \sum_{k=1}^{K} \beta_{11}(k)\Delta P_{t-k} + \sum_{k=1}^{K} \beta_{12}(k)\Delta X_{t-k} + \Omega_t \end{aligned} \tag{3}$$

where t refers to time (t = 1, 2, …, 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 one, 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.

Price relationships among market locations can be 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:

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

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 prices of refined, bleached, and deodorized (RBD) soybean oil 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. 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). Lag length for the entire period is set to 2, lag length is set to 4 for the period before the biodiesel surge, and lag length is set to 1 for the period after the biodiesel surge.

Results

Determining market integration starts with the ADF tests on the price series. Table 2 shows that, based on the ADF tests, the data do not provide enough evidence to reject the null hypothesis of a unit root's existence. The price series are, therefore, non-stationary. However, long-run relationships among the price series can still be estimated as long as each one is integrated of the same degree. The same table 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.

In Table 3, we extend the ADF tests to analyze the existence of unit roots before and after the U.S. biodiesel production increase during the mid-2000s. Overall, the price data are also found to be integrated of order 1 in both time periods. It should be noted, however, that the tests provided stronger evidence for rejecting the null hypothesis of the existence of unit roots after the biodiesel production increase. This may provide evidence that the biodiesel production increase has caused structural changes in these prices.

We next determine whether the price series are 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 4 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. 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 also reported in Table 4. Two observations should be noted. First, the price series exhibit cointegrating relationships before and after to the sudden U.S. biodiesel production increase. Second, the significance level in rejecting the null hypothesis of no cointegration is lower after the sudden U.S. biodiesel production increase for several market pairings. One implication from this particular observation is that the sudden U.S. biodiesel production increase may have had weakened soybean oil spatial price relationships.

Following Pendell and Schroeder (2006), VAR model (3) was next estimated to test the strength of the price linkage 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). Except for two market pairings, the results lend support to full integration among these markets, either before or after the sudden U.S. biodiesel production increase (Table 5). 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 also after the sudden increase. The exception is West Coast/East Coast and the East Coast/West Coast market pairings. Data do not provide evidence that the prices in these markets are not fully integrated after the biodiesel production increase. This provides further evidence that the biodiesel production increase has weakened price relationships. For the other market pairings, the VAR model's results indicate that soybean oil prices were fully integrated both before the U.S. biodiesel production increase and after the increase.[5]

Speed-of-adjustment coefficients from the VEC model (4) are reported in Table 6. 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. Take 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 0.43 units, and the U.S. Gulf price rises by 0.29 units. Both changes are larger than the degree of adjustment before the marked biodiesel production increase (–0.27 and –0.07, 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.

[5] It should be noted, however, that while statistical evidence still show the Central Illinois/East Coast and East Coast/Central Illinois market price pairings are fully integrated after the increase in biodiesel production, the VEC results show a substantial increase in level of significance, bordering on rejection of the null hypothesis. This may again indicate that the biodiesel production increase have weakened price relationships.

The sudden increase in U.S. biodiesel production's effect is also evident when looking at Granger causality tests. The Granger causality test is another useful tool in analyzing the price relationships among the four markets because it examines whether the future prices from one market can consistently be better predicted using historical prices from another market (Granger, 1969). Tables 7 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 three market pairings show two-way Granger causality in prices before the U.S. biodiesel production increase. The exceptions are prices in the West Coast market, which Granger cause prices in the West Coast and East Coast markets, and prices in the East Coast Market, which Granger causes prices in the West Coast market. All but two Granger causalities cease to exist, however, after the production increase. While weakened, prices in the U.S. Gulf market sill Granger-causes prices in the West Coast and East Coast markets. 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.

Conclusions

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 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 has 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 transactions costs (Barrett, 1996; Goodwin and Piggott, 2001). Balke and Fomby (1997) discussed that the tendency for two variables to move toward a long-run equilibrium 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 et al., 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 above, 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 US agricultural commodity markets. Among those is Yang and Leatham (1998), which tests the market efficiency of the US 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 aren’t available to consider these relationships. Two factors in particular are worth noting. First, 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 2000/2001 to 92.27 million metric tons in 2011/2012 (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.

[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).

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Table 1. Soybean Oil Price: Summary Statistics (Cents per Pound)

Market	Mean	Std. Dev.	Min.	Max.
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. Test Statistics for Unit Root on Soybean Oil, Entire Period

	Levels		Differenced	
Market	Augmented Dickey-Fuller	Zivot-Andrews	Augmented Dickey-Fuller	Zivot-Andrews
Central Illinois	– 1.89	– 4.28	– 4.20 *	– 7.17 *
U.S. Gulf	– 1.83	– 4.35	– 4.03 *	– 7.02 *
West Coast	– 1.82	– 4.39	– 4.27 *	– 7.13 *
East Coast	– 1.84	– 4.46	– 4.28 *	– 7.15 *

Notes: The null hypothesis is that a unit root exists, i.e. time series is not stationary. Augmented Dickey-Fuller tests are performed with two lags based on Akaike Information Criterion. * denotes statistical significance at the 1% level.

Table 3. Augmented Dickey-Fuller Test for Unit Root on Soybean Oil, Before and After Production Surge in Biodiesel

	Before Biodiesel				After Biodiesel			
	Levels		Differenced		Levels		Differenced	
Market	ADF-stat	P-value	ADF-stat	P-value	ADF-stat	P-value	ADF-stat	P-value
Central Illinois	– 0.43	0.90	– 2.94	0.04	– 2.30	0.17	– 3.58	0.01
U.S. Gulf	0.73	0.99	– 2.58	0.10	– 2.26	0.18	– 3.46	0.01
West Coast	– 0.02	0.96	– 2.43	0.13	– 2.37	0.15	– 3.69	0.00
East Coast	– 0.08	0.95	– 2.56	0.10	– 2.39	0.14	– 3.69	0.00

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

Table 4. Results from Johansen Cointegration Tests

	Trace Statistics		
Market Pairs	Entire Period	Before Biodiesel	After Biodiesel
Central Illinois/U.S. Gulf	2.81 *	3.26 *	2.43 *
Central Illinois/West Coast	3.09 *	0.21 *	2.32 *
Central Illinois/East Coast	3.17 *	0.00 *	2.30 *
U.S. Gulf/West Coast	2.66 *	4.77 **	2.07 *
U.S. Gulf/East Coast	2.66 *	4.07 **	2.08 *
West Coast/East Coast	2.86 *	1.21 *	2.94 *

Notes: Critical values are 6.65, 3.76 for the 1%, 5% level of significance, respectively. **, * denotes significance at the 1%, 5%, respectively. Using Akaike Information Criterion, lag length for the entire period is set to 2, lag length is set to 4 for the period before the biodiesel surge, and lag length is set to 1 for the period after the biodiesel surge. Entire period consists of 96 monthly observations. Number of samples is 25 prior to production surge in biodiesel (May 2007), and 71 after the surge.

Table 5. VAR Parameter Estimates from Regime Shift Model

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

Notes: **, * significant 1%, 10%, respectively. Standard errors in parenthesis. Lag length is set to 2. Number of samples is 92.

Table 6. Speed-of-Adjustment Coefficients from Vector Error Correction Model

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 (035)	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

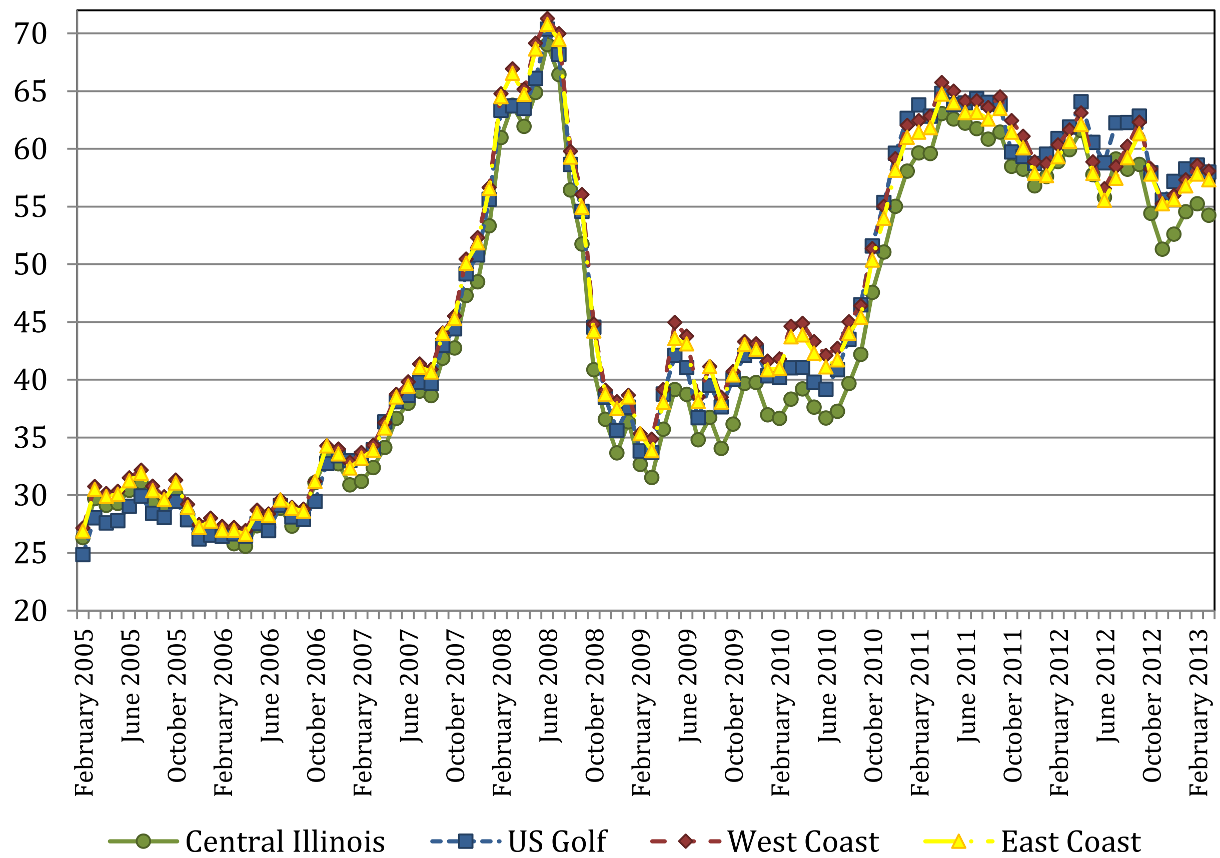
Notes: Net impact may not add up due to rounding. Standard errors in parenthesis. Lag length is set to 2. Number of samples is 95.

Table 7. Granger Causality for Soybean Oil Prices from Vector Error Correction Model

Dependent Market/ Independent Market	χ^2 Test Statistic		
	Pre-Biodiesel	Post-Biodiesel	Entire Period
Central Illinois/U.S. Gulf	30.41 **	1.99	1.07
Central Illinois/West Coast	6.00	0.01	1.41
Central Illinois/East Coast	10.16 *	0.00	1.15
U.S. Gulf/Central Illinois	20.44 **	0.49	0.97
U.S. Gulf/West Coast	35.35 **	0.65	0.87
U.S. Gulf/East Coast	24.67 **	1.03	0.77
West Coast/Central Illinois	10.73 *	1.76	4.48
West Coast/U.S. Gulf	28.72 **	3.83 *	3.66
West Coast/East Coast	5.66	0.20	0.46
East Coast/Central Illinois	14.15 **	1.88	4.03
East Coast/U.S. Gulf	24.83 **	4.30 *	3.58
East Coast/West Coast	5.15	0.47	0.71

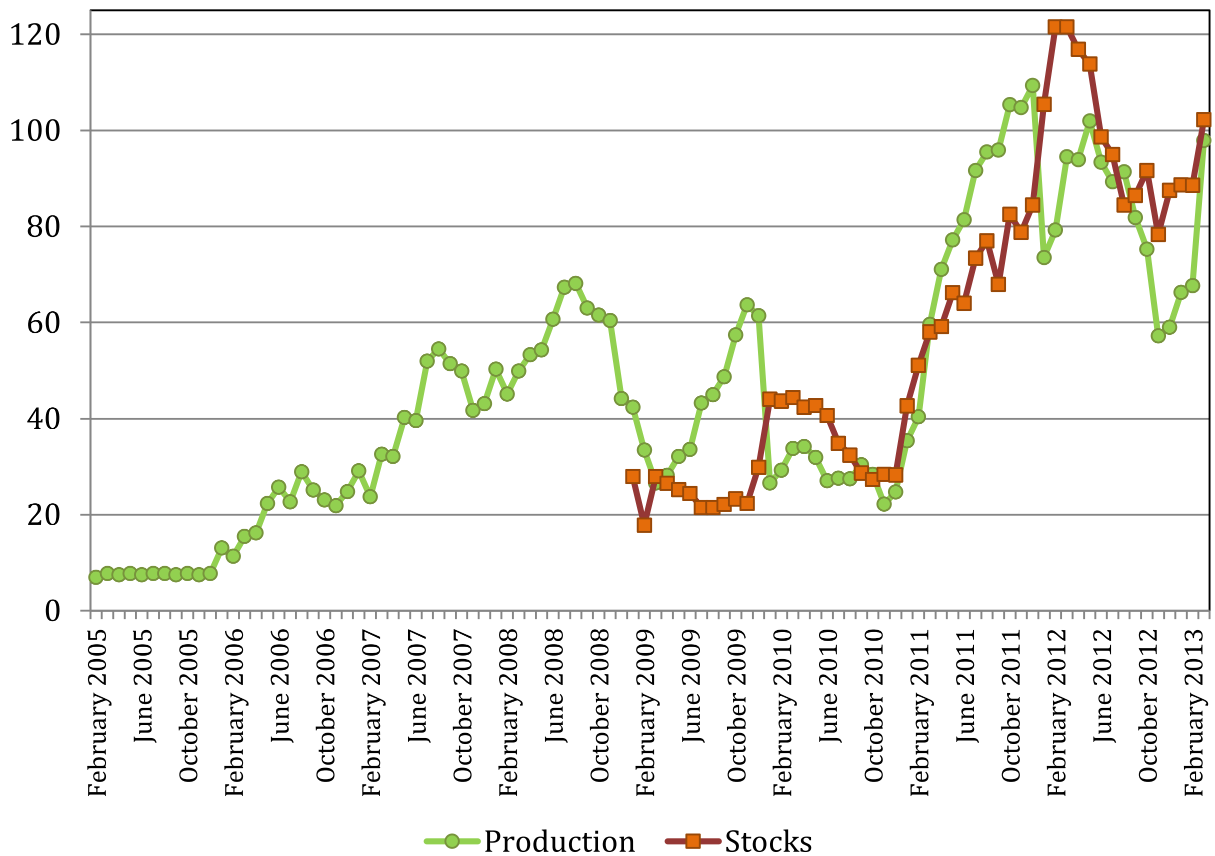
Note: **,* significant 1%, 5%, respectively. Using Akaike Information Criterion, lag length for the entire period is set to 2, lag length is set to 4 for the period before the biodiesel surge, and lag length is set to 1 for the period after the biodiesel surge.

Figure 1. Monthly Average Refined-Bleached-Deodorized Soybean Oil Prices (cents per pound) from February 2005 to March 2013



Source: The Jacobsen

Figure 2. Monthly U.S. Biodiesel Production and Stocks (million gallons) from January 2001 to September 2013



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