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Analyzing Factors Affecting U.S. Food Price Inflation.

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ABSTRACT

Since the summer of 2007, U.S. food price has increased dramatically. Given public anxiety over fast-rising food prices in recent years, this paper attempts to analyze the effects of market factors ─ prices of energy and agricultural commodities and exchange rate ─ on U.S. food prices using a co-integration analysis. Results show that the agricultural commodity price and exchange rate play key roles in determining the short- and long-run movement of U.S. food prices. It is also found that in recent years, the energy price has been a significant factor affecting U.S. food prices in the long-run, but has little effect in the short-run. This implies the strong long-run linkage between energy and agricultural markets has emerged through production of commodity-based ethanol in the recent years.

Keywords: Agricultural commodity price, Energy price, Exchange rate, Food price inflation, Time-series analysis

HIGHLIGHTS

Since the summer of 2007, U.S. food price has increased dramatically as U.S. consumers have begun to face rising food prices at supermarket checkout lines. During the second half of 2007, the food CPI increased by approximately 4.5%, the largest annual jump in 17 years. According to the U.S. Department of Agriculture (USDA), the food CPI increased to 5.5% in 2008. Although the factors driving up the recent food price inflation are complex, higher farm commodity prices, energy prices and weakening U.S. dollar are considered to be the major culprits behind the rapid spike in U.S. food prices.

The main objective of this study is to examine the dynamic effects of market factors on U.S. food prices in a co-integration framework. The empirical focus is on the assessment of the short- and long-run linkages between changes in U.S. food prices and changes in prices of energy and agricultural commodities and exchange rate. For this purpose, the Johansen co-integration and a vector error-correction (VEC) model are used.

The results show that the agricultural commodity prices and exchange rate play crucial roles in influencing the short- and long-run behavior of U.S. food prices. It is also found that the energy price has been an important factor affecting U.S. food prices in recent years in the long-run, but has a little impact in the short-run. This further suggests that, in the long-run, the linkage between energy and agricultural commodity/food markets has been quite strong, due mainly to crop-based bio-fuel production. The energy market is also closely related to the movement of the value of the U.S. dollar against its major trading partners. Under the given linkage between the agricultural commodity prices and exchange rate which is linked to the energy market, the uncertainty on commodity prices could increase as the value of the U.S. dollar fluctuates against the currencies of its major trading countries.

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INTRODUCTION

U.S. food price inflation has been relatively low and stable over the past 15 years. From 1991 through 2006, the Consumer Price Index for all food (food CPI) has risen by approximately 2.5% annually (Figure 1). Since the summer of 2007, however, this trend changed dramatically as U.S. consumers have begun to face rising food prices at supermarket checkout lines. During the second half of 2007, the food CPI increased by approximately 4.5%, the largest annual jump in 17 years. According to the U.S. Department of Agriculture (USDA), the food CPI increased to 5.5% in 2008. The objective of this paper is to examine the short- and long-run linkages between changes in U.S. food prices and changes in market factors such as prices of energy and agricultural commodities, ethanol production and exchange rate.

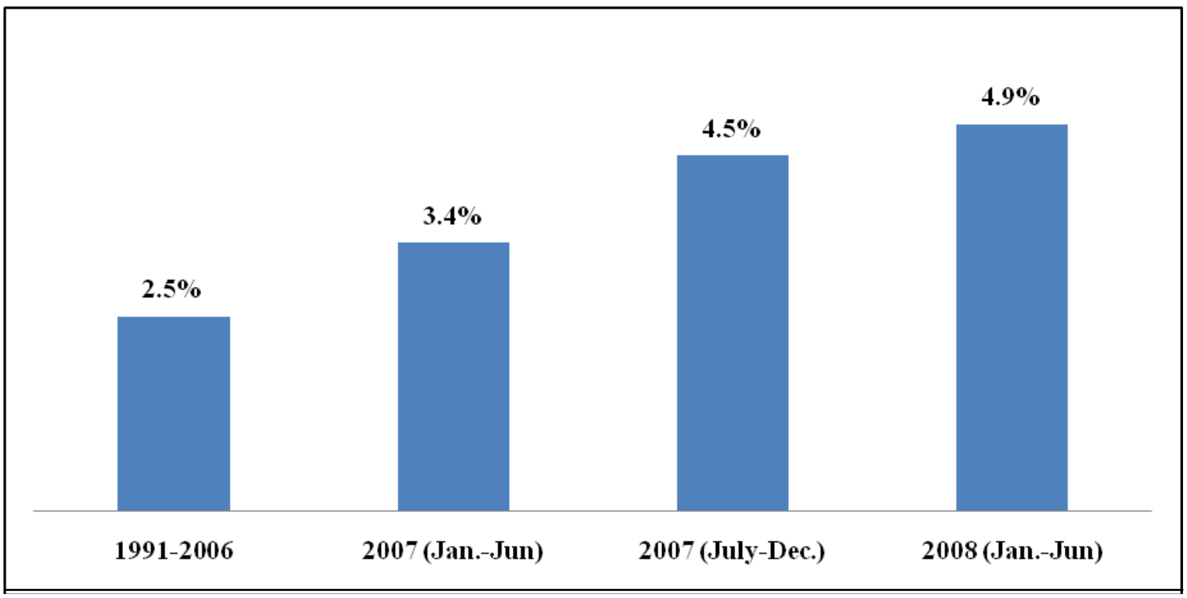

Figure 1. Consumer Price Index for all food (food CPI) in the United States

A number of studies have tackled the recent surge in U.S. food prices over the last two years (see Abbott et al. (2008) for detailed overview of these studies). Some studies done by national and international institutions such as the International Food Policy Research Institute (IFPRI), U.S. Department of Agriculture (USDA) and Congressional Research Service (CRS) have typically focused on identifying the causes and sources of current food price increases. Other studies financed or produced by interest groups such as the American Coalition for Ethanol (ACE), Grocery Manufacturer's Association (GMA) and National Corn Growers Association (NCGA) have been mainly conducted to support position of those organizations in the most favorable light possible. For example, the NCGA released two reports, "Understanding the Impact of Higher Corn Prices on Consumer Food Prices" and "U.S. Corn Growers: Producing Food and Fuel," in 2007, claiming that corn-based ethanol production is not the primary driver of rising food prices. In all these studies, however, their assessments have been conducted on the basis of descriptive statistics and graphical methods that mainly use U.S. crop production and price data. Until recently, no study has relied on an econometric technique in investigating factors affecting fast-rising food prices in the United States. This paper, therefore, attempts to investigate the issue using the Johansen co-integration test and a vector error-correction (VEC) model. The remaining sections present background information, empirical methodology, empirical findings, and conclusions.

FACTORS AFFECTING U.S. FOOD PRICE INFLATION

Although the factors driving up the recent food price inflation are complex, higher farm commodity prices and energy prices are considered to be the major culprits behind the rapid spike in U.S. food prices. Specifically, significant growth in the use of farm commodities (i.e., corn) for bio-fuel production and increased U.S. exports driven by a weak dollar result in record or near-record prices for key commodities such as corn, soybeans and wheat in 2007. With double-digit growth rates (24.5% annually) since 2002, production of fuel ethanol reached a record high of 9.2 billion gallons in 2008, a 41.6% increase over 2007 (Figure 2a). Accordingly, in the 2007/08 marketing year, over 3 billion bushels of corn (23% of the harvested corn) were used to produce ethanol, more than a 30 % increase from the previous year. Corn prices in turn climbed up to $4.78 per bushel in 2008, an approximately 43% increase from the average of the previous 15 years ($2.37) (Figure 2b).

Corn competes with soybeans and wheat for cropland thus higher corn prices have motivated farmers to increase corn acreage at the expense of soybeans and wheat, contributing to tight supplies and higher prices for those crops. Accordingly, soybean and wheat prices rose from average prices of $5.85 per bushel and $3.38 per bushel during 1991-2006 to $11.30 per bushel and $8.03 per bushel in 2008, respectively (Figure 2b).

Another factor is a significant shortfall in agricultural production in major agricultural exporting countries. Australia, for example, produced only 10.5 and 13.1 million metric tons of wheat in 2006 and 2007, compared to a normal 21 and 24 million metric tons. Other countries such as Argentina also experienced a significant decrease in wheat production in 2006 and 2007, down as much as 11% of the 2000 production.

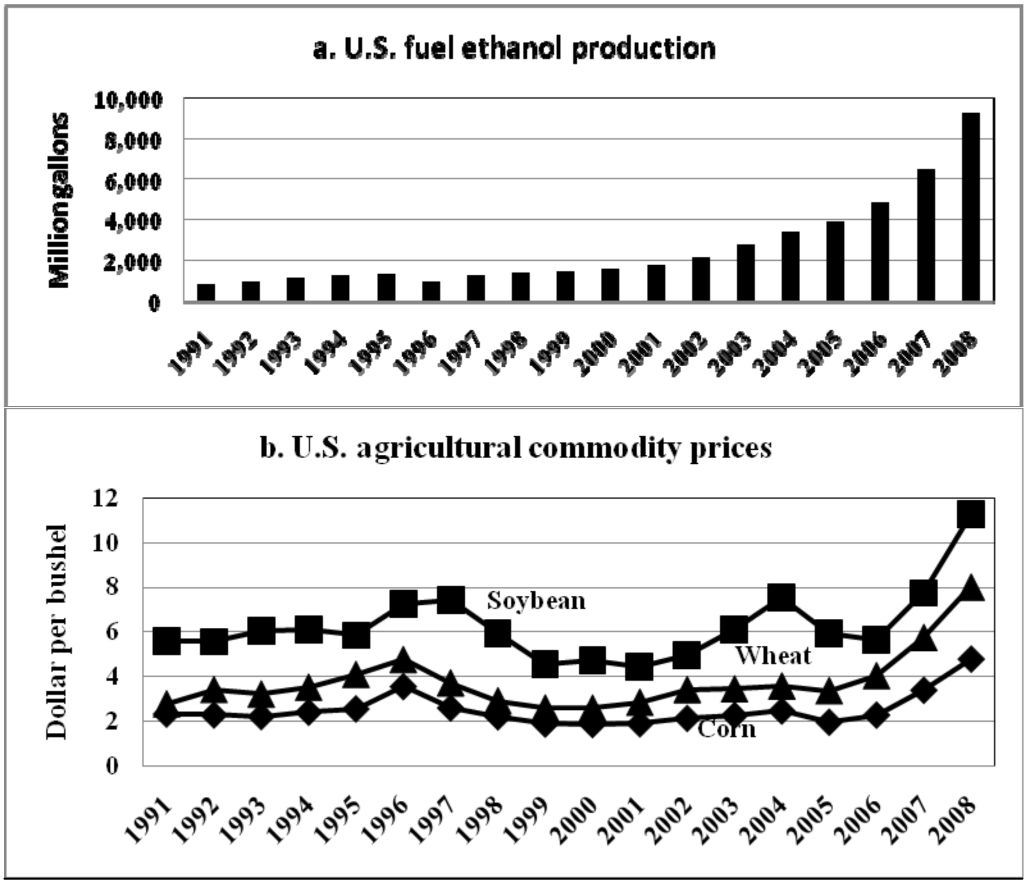

Figure 2. U.S. fuel ethanol production and agricultural commodity prices, 1991-2008

In addition, the U.S. dollar's global weakness over the past five years has helped make U.S. agricultural commodities more competitive in the world market, thereby boosting demand for U.S. agricultural commodities and prices (Figure 3a). Between 2002 and May 2008, the value of the U.S. dollar has weakened against its major trading partners such as Canada (33.5%), Japan (14.0%), Korea (14.2%) and Taiwan (7.1%). The U.S. dollar also has declined against its main trade competitors such as the EU (38.2%), Brazil (38.0%) and China (14.2%) during the same period (Figure 3b).

Higher domestic commodity prices tend to decrease commodity exports, but this weak dollar has offset that impact and made U.S. commodity more attractive in the international market. As a result, U.S. corn exports are expected to reach at a record 2.45 billion bushels in the 2007/08 market year. Soybean exports are projected to increase by nearly 40% of U.S. soybean crops in 2007, up from 35% the year before. Combining rapidly escalating demand for bio-fuel production and a weak dollar, commodity prices are expected to increase substantially in the future. According to the USDA's projections, the prices of corn, soybeans and wheat will go up to \$5.30-\$6.30 per bushel, \$11.00-\$12.50 per bushel and \$6.75-\$8.25 per bushel, respectively, in the 2008/09 marketing year.

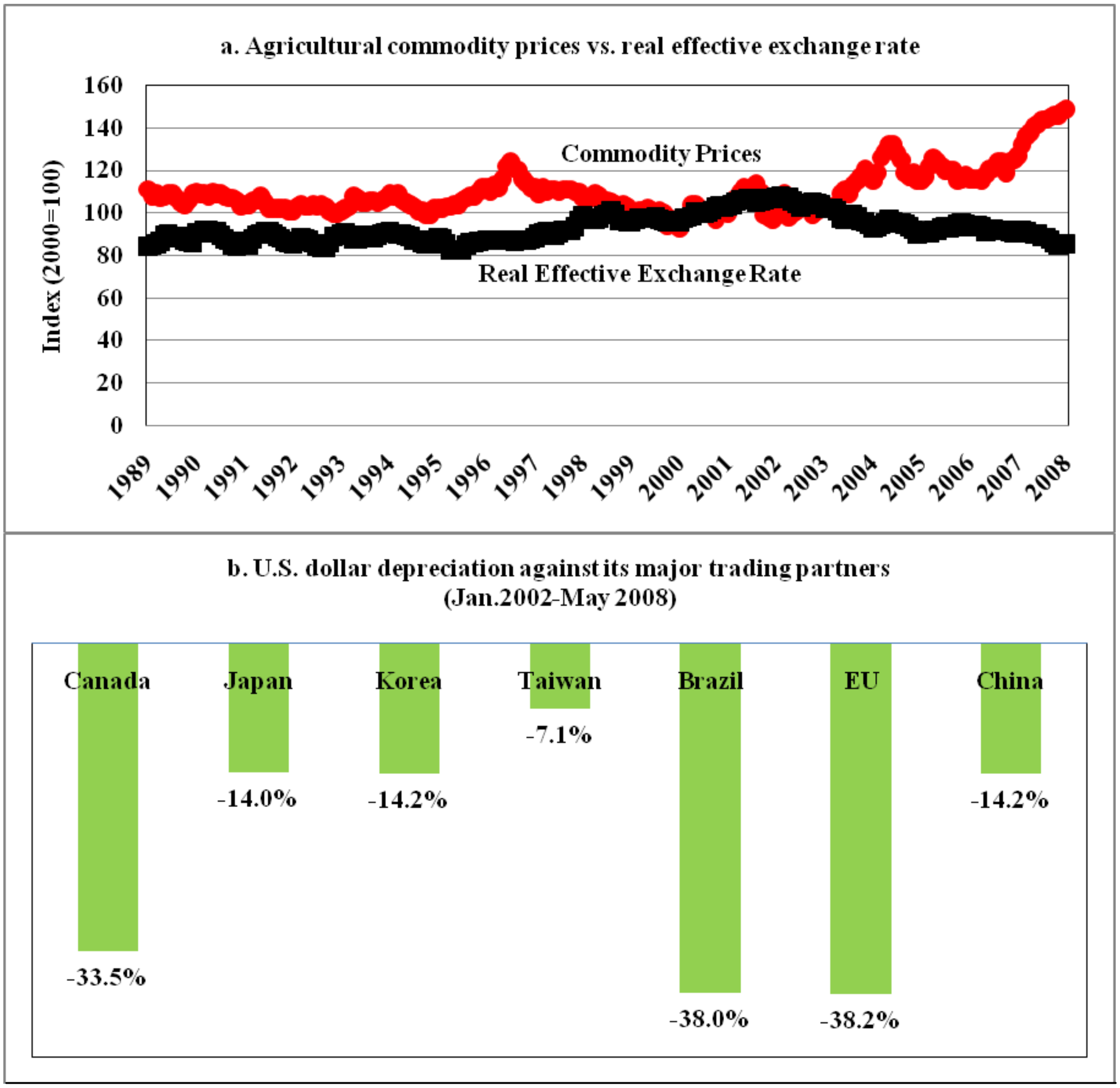

Figure 3. U.S. agricultural commodity prices and real effective exchange rate, 1989-2008

High energy prices have also contributed to the recent drastic surge in U.S. food prices. For example, the price of crude oil was $27 per barrel in December 2003, but it went up to $62 per barrel in April 2006, then to $98 per barrel in March 2008, and more than $120 per barrel in June 2008 (Figure 4). The rapid increase in crude oil prices has significantly raised the costs of producing and shipping agricultural commodities through increases in the prices of fertilizer, diesel, agricultural chemicals and other inputs. In addition, high crude oil price means high gasoline price, which increases production of ethanol, further boosting demand for corn and thus prices (Figures 2a and b).

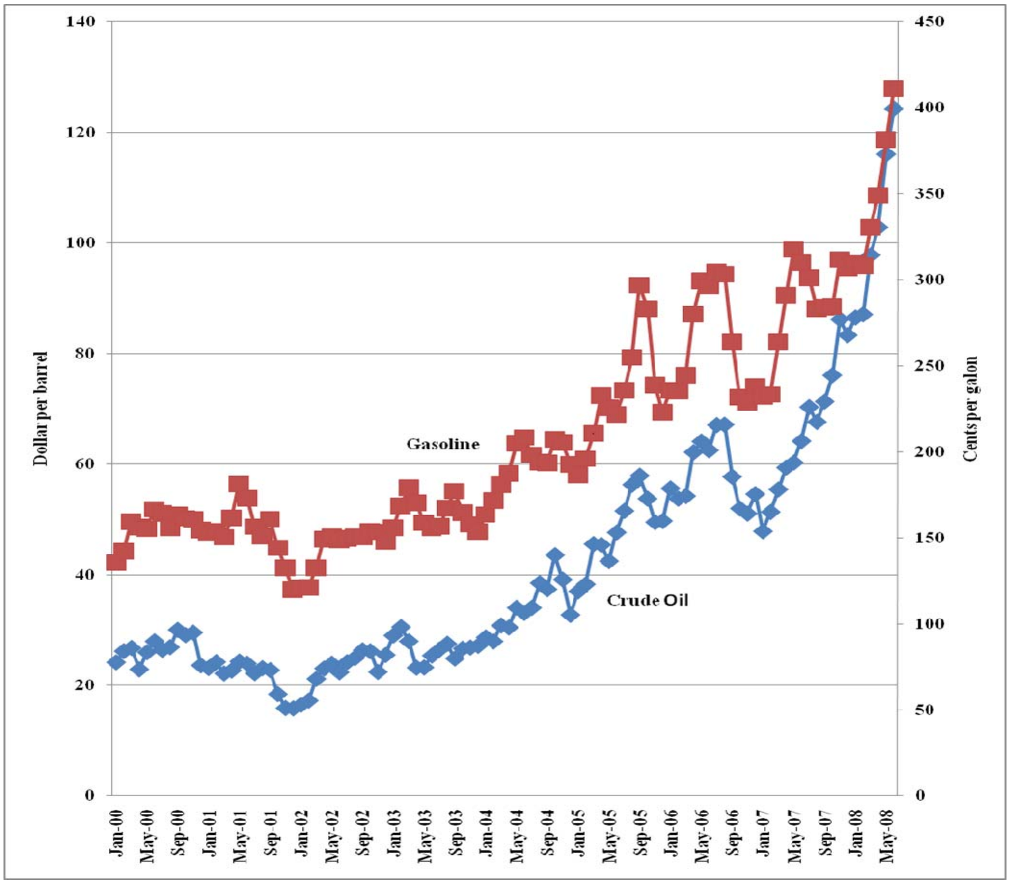

Figure 4. Crude oil and gasoline prices in the United States, Jan. 2000-May 2008

DEVELOPMENT OF AN EMPIRICAL MODEL

The Johansen multivariate co-integration analysis and a vector error-correction (VEC) model are employed for empirical analysis (see Appendix 1 for details on empirical methodology). The Johansen approach is used to identify the long-run equilibrium relationships (co-integration) among the selected variables, (food price, commodity price, energy price, exchange rate, and ethanol production) while the VEC model provides information on the short-run dynamic adjustment to changes in the variable with the model. If the variables are co-integrated, they tend to move closely together over time; hence, a test for co-integration is identical to a test of a long-run equilibrium to which an economic system converges over time. In addition, structural break(s) are incorporated into the analysis because failure to account for structural break(s) in co-integration analysis may raise issue of spurious long-run relationships (Harris and Sollis 2003).

Data

Since the main focus of this paper is on the assessment of factors affecting changes in U.S. food prices (FP_t), as noted earlier, market variables that are thought to be of central importance in influencing U.S. food prices are selected for inclusion in the model. These variables are agricultural commodity prices (CP_t), energy prices (EP_t), ethanol production (ETH_t) and exchange rates (ER_t).

The U.S. consumer price index for all food (2000=100) is used as a proxy for U.S. food prices and is collected from the Bureau of Labor Statistics (BLS) in the U.S. Department of Labor (USDOL). The prices received index for all farm products (2000=100) is used as a proxy for U.S. commodity prices and is obtained from the National Agricultural Statistics Service (NASS) in the U.S. Department of Agriculture (USDA). The U.S. index for energy is used as proxy for U.S. energy prices and is taken from the BLS. The fuel ethanol production (measured in million gallons) is used as a proxy for crop-based bio-fuel production and is collected from the Monthly Energy Review (MER) published by the Energy Information Administration (EIA). Finally, the exchange rate is a real effective exchange rate index (2000=100) and is collected from the International Financial Statistics (IFS) published by the International Monetary Fund (IMF). Since the real effective exchange rate is defined as the currencies of trading partners per unit of the U.S. dollar, a decline in exchange rate indicates a real depreciation of the U.S. dollar. Monthly seasonally adjusted data are collected for the period from January 1989 to January 2008. All variables are converted to natural logarithms and used throughout.

Unit Roots under Structural Shifts

When dealing with time-series data, the possibility of non-stationarity (or unit roots) in a series raises issues about parameter inference and spurious regression (Wooldridge 2000). A stationary series is defined as a series that tends to return to its mean value and fluctuate around it within a more or less constant range. A non-stationary series (unit roots), on the other hand, is defined as a series that has a different mean at different points in time and its variance changes with the sample size (Harris and Sollis 2003). The OLS regression involving non-stationary series no longer provides the valid interpretations of the standard statistics such as t-statistics, F-statistics and confidence intervals. To address this problem, non-stationary variables should be differentiated to make them stationary. Engle and Granger (1987), however, prove that, even in the case that all the variables in a model are non-stationary, it is possible for a linear combination of integrated variables to be stationary; in this case, the variables are said to be co-integrated and the problem of spurious regression does not arise. Hence, testing for stationarity (non-stationarity) in time-series data is a prerequisite for the application of the co-integration technique.

In addition, since structural change is an important issue in time-series analysis and affects inference on unit roots and co-integration, it is important to allow for its possibility at the estimation process (Maddala and Kim 1998). Recent studies, for example, show that one major drawback of the standard unit root tests (e.g., augmented Dickey-Fuller (ADF) test) is that, in all of them, the implicit assumption is that the deterministic trend is correctly specified (Perron 1989, Banerjee et al. 1992, Maddala and Kim 1998). In other words, if there is a break in the deterministic trend, the standard unit root tests could falsely lead to the conclusion that there is a unit root when in fact there is not (Perron 1989). Moreover, failure to account for structural breaks in co-integration analysis may raise the issue of spurious long-run relationships (Harris and Sollis 2003). Hence, it is both desirable and necessary to perform tests for structural breaks of the series to overcome the weaknesses of the standard unit root procedure, as well as to obtain improved econometric estimates of long-run relationships in the series, which we have done in this section.

Before conducting a unit root test that allows for the possibility of structural break(s), we first present graphical inspection for the five series. The graph of the U.S. food price, for example, shows that there is a change in the intercept of the series in the late 1998 (Figure 5a).

The graph of the energy price, on the other hand, shows that there are changes in both the intercept of the series in early 1998 and in the slope afterwards (Figure 5b). The same feature appears to hold for the commodity price (in late 1998), exchange rate (in late 2001) and ethanol production (in early 1996). To verify this graphical inspection, we follow the Perron (1989) testing procedure (see Appendix 2 for details). The results show that the identified breakpoints (T_B) are December 1998 for the U.S. food price, November 1998 for the commodity price, February 1998 for the energy price, October 2001 for the exchange rate, and February 1996 for the ethanol production.

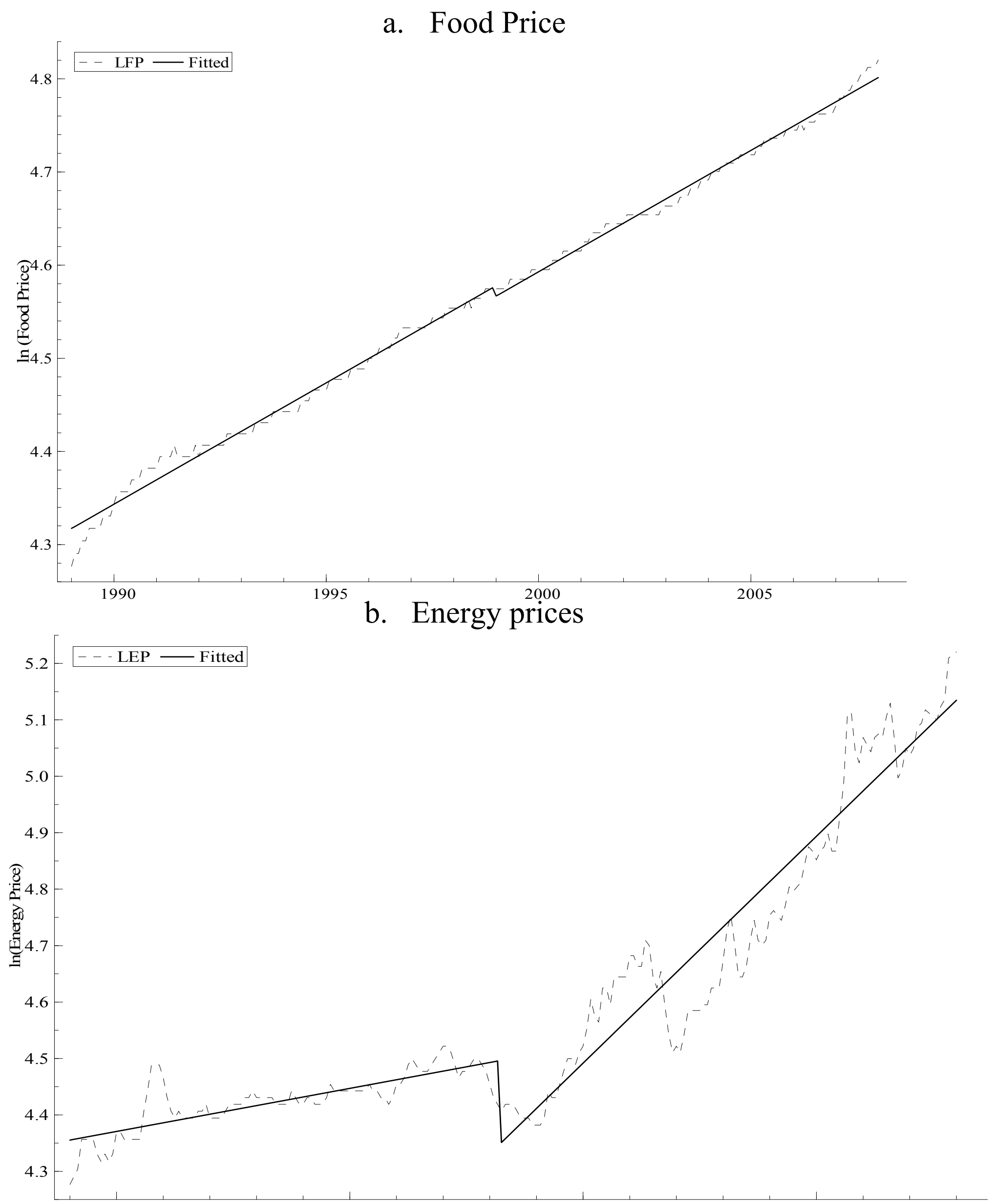

Figure 5. Logarithm of U.S. food and energy prices, actual values and modeled structural shifts (fitted values)

Having obtained the identified breakpoints, we conduct unit root tests that allow a shift in the intercept and that in the intercept and trend developed by Perron (1989) (see Appendix 2 for details). The results show that, even when structural shifts are included, the null hypothesis of non-stationary cannot be rejected at the 5% level for four of them: the food price, commodity price, energy price and exchange rate (first column of Table 1). However, the null hypothesis can be rejected for the ethanol production even at the 10% level. The results thus lead to conclusion that the underlying process for the U.S. ethanol production can be characterized by stationary fluctuations around a deterministic trend function.

Given the results of the Perron's testing procedure, it is no longer appropriate to use the full sample that includes stationary ethanol production series in the co-integration analysis. As an alternative, therefore, the full sample is divided into two subsamples according to the most recent break point (pre- and post- November 2001) in order to see if this feature is stable in both cases. Subsample I covers January 1989-October 1998 and subsample Π covers November 2001-January 2008.[1] With these two subsamples, we conduct unit root tests using the Dickey-Fuller generalized least squares (DF-GLS) test (Elliott et al. 1996). The DF-GLS test statistics in both subsamples are estimated from a model that includes a constant and a trend variable. The results show that, with subsample I, the null hypothesis of non-stationarity cannot be rejected at the 5% level for all the series except the ethanol production (second column of Table 1). With subsample Π, on the other hand, the null hypothesis cannot be rejected for all the five series (third column of Table 1). From these findings, we conclude that the food price, commodity price, energy price and exchange rate in subsamples I and Π are non-stationary. However, since the ethanol production in subsample I is found to be stationary, it cannot be used for the co-integration analysis. For further time-series analysis, therefore, this paper focuses on the four variables ─ FP_t, CP_t, EP_t, and ER_t.

Table 1. Results of unit root tests

	Perron test	DF-GLS test	
	Full sample (Jan. 1989-Jan. 2008)	Subsample I (Jan. 1989-Oct. 1998)	Subsample Π (Nov. 2001-Jan. 2008)
$\ln FP_t$	-3.02 (6)	-1.18 (2)	-1.11 (6)
$\ln CP_t$	-3.62 (1)	-1.86 (1)	-2.07 (1)
$\ln EP_t$	-3.25 (2)	-1.35 (7)	-1.77 (2)
$\ln ER_t$	-2.59 (6)	-2.42 (4)	-1.02 (1)
$\ln ETH_t$	-5.22 (3)**	-3.45 (3)**	-0.57 (12)

Note: ** denotes rejection of the null hypothesis at the 5% levels. Lag lengths are given inparentheses. The 5% and 10% critical values for the Perron test are -3.76 and -3.46 for FP_t, -4.24 and -3.96 for CP_t and EP_t, -4.18 and -3.86 for ER_t, and -4.22 and -3.95 for ETH_t, respectively, which are obtained from Tables 4B and 6B in Perron (1989). The 5% and 10% critical values for the DF-GLS including a constant and a trend are – 3.01 and -2.72 for Subsample I and -3.02 and -2.73 for Subsample Π, respectively. The lag order for the Perron test is chosen by a test on the significance of the estimated coefficients of lagged first-differences, while that for the DF-GLS is selected by the Schwert criterion (SC).

[1] Given that the breakpoints found here vary for each series as they go from November 1998 to October 2001, excluding these breakpoints for constructing sub-samples allows us to estimate the maintained hypothesis more properly and accurately ─ that the U.S. food price and its major determinants are much closely linked in recent years. For this reason, we have eliminated the period of these breakpoints (November 1998-October 2001) for further analysis.

EMPIRICAL RESULTS

Analyzing Long-Run Relationships

With the selected lag lengths (k=5 for subsample I and k=6 for subsample II) in non-stationary VAR models, the Johansen co-integration procedure is conducted to determine the number of co-integrating relationships among the four variables.[2] The results of co-integration estimation show that two co-integration vectors ($r = 2$) for subsample II at the 5% level, whereas no co-integration is found for subsample I (Table 2). Specifically, with subsample II, the trace tests show that the null hypotheses of no co-integration ($r = 0$) and one co-integration ($r = 1$) can be rejected, but fail to reject the null of two co-integration vectors ($r = 2$). With subsample I, on the other hand, the lower trace statistic fails to reject no co-integration even at the 10% level, indicating that the four variables are not co-integrated over the 1989-1998 period. These tests thus provide evidence to support the hypothesis that the U.S. food price and its major determinants such as the commodity price, energy price and exchange rate have been much more closely linked and have the long-run relationships among them in the 2001-2008 period compared to the early period. Note that the VAR model includes an unrestricted constant and a linear trend in both subsamples.

Table 2. Results of Johansen co-integration rank tests

Subsample I (Jan. 1989-Oct. 1998)		
Null hypothesis	Eigenvalue	Trace statistics
H_0: r = 0	0.2123	59.63 [0.11]
H_0: r ≤ 1	0.1389	32.67 [0.36]
H_0: r ≤ 2	0.0792	15.77 [0.52]
H_0: r ≤ 3	0.0554	6.44 [0.42]
Subsample Π (Nov. 2001-Jan. 2008)		
Null hypothesis	Eigenvalue	Trace statistics
H_0: r = 0	0.3666	63.11 [0.00]**
H_0: r ≤ 1	0.2160	31.59 [0.03]**
H_0: r ≤ 2	0.1415	11.79 [0.11]
H_0: r ≤ 3	0.0261	1.82 [0.18]
Full sample (Jan. 1989-Jan. 2008)		
Null hypothesis	Eigenvalue	Trace statistics
H_0: r = 0	0.2096	91.37 [0.00]**
H_0: r ≤ 1	0.0862	38.68 [0.13]
H_0: r ≤ 2	0.0556	18.49 [0.32]
H_0: r ≤ 3	0.0249	5.67 [0.51]

Note: ** denotes rejection of the null hypothesis at the 5% significance level. p -values are given in parentheses.

[2] Diagnostic tests on the residuals of each equation and corresponding vector test statistics support the VAR models with five lags for subsample I and six lags for subsample II. The test results are not reported here for brevity but can be obtained from the authors on request.

Having obtained two co-integrating vectors in subsample II, the test for the long-run weak exogeneity is conducted to examine the absence of long-run levels of feedback due to exogeneity (Johansen and Juselius 1992). A weakly exogenous variable can be interpreted to be a driving variable, which pushes the other variables in adjusting to the long-run equilibrium, but is not influenced by the other variables in the model. The long-run weak exogeneity test can be done by restricting parameter in speed-of-adjustment (α) to zero in the model. The results show that, with subsample II in which the four variables are co-integrated, the null hypothesis of weak exogeneity cannot be rejected for the energy price and exchange rate even at the 10% level (Table 3), indicating that these two variables are weakly exogenous to the long-run relationships in the model. This finding further suggests that the energy price and exchange rate are the driving variables in the system and significantly influence the long-run movements of the U.S. food and commodity prices, but are not affected by the U.S. food and commodity prices.

Table 3. Results of weak exogeneity tests

Subsample Π (Nov. 2001-Jan. 2008)	
Variable	Weak exogeneity $H_0: \alpha_i = 0$
FP_t	9.47 [0.00]**
CP_t	5.57 [0.06]*
EP_t	3.89 [0.14]
ER_t	2.74 [0.26]
Full sample (Jan. 1989-Jan. 2008)	
Variable	Weak exogeneity $H_0: \alpha_i = 0$
FP_t	29.09 [0.00]**
CP_t	1.79 [0.18]
EP_t	1.92 [0.18]
ER_t	0.65 [0.42]

Notes: FP_t, CP_t, EP_t and ER_t represent U.S. food price, U.S. commodity price, U.S. energy price and exchange rate, respectively. α_i represents the speed of adjustment to equilibrium. LR test statistic is based on the χ^2 distribution and parentheses are p-values. ** and * denote the rejection of the null hypothesis at the 5% and 10% significance levels, respectively.

The co-integration vectors (β_i) is used to describe the long-run relationship among the variables.[3] With two co-integrating vectors in subsample II, for example, the two error-correction models ($ec1(\hat{\beta}_1)$ and $ec2(\hat{\beta}_2)$) can be expressed in a reduced-form as follows:

$$ec_1 : FP_t = \underset{(3.65)}{0.114}EP_t + \underset{(6.34)}{0.446}CP_t + \underset{(2.35)}{0.001}trend \qquad (1)$$

$$ec_2 : CP_t = \underset{(2.44)}{0.355}EP_t - \underset{(4.47)}{4.531}ER_t + \underset{(3.32)}{0.002}trend \qquad (2)$$

Equations (1) and (2) show that all estimates are statistically significant at the 5% level (parentheses are t-values) and have the expected signs. Specifically, the U.S. food price has a positive long-run relationship with the energy, and commodity prices. This suggests that an increase in energy price leads to a rise in food price through the increased costs of producing and shipping agricultural commodities. Additionally, the U.S. food price has a positive long-run relationship with the commodity price, suggesting that an increase in commodity price causes a rise in the food price in the long-run. The commodity price has a positive long-run relationship with the energy price, indicating that an increase in energy price leads to an increase in input costs in producing agricultural commodities, thereby resulting in a rise in commodity prices. Furthermore, an increase in energy price raises the price of corn through increased use of corn for corn based ethanol production. Finally, the commodity price has a negative long-run relationship with the exchange rate. This indicates that the depreciation of the U.S. dollar indeed makes U.S. agricultural commodities more competitive in the world market, thereby boosting demand for U.S. agricultural commodities and hence prices. Notice that equations (1) and (2) show no direct relationship between the exchange rate and the U.S. food price in the long-run.

Since one of the key motivations of this study is to examine the long-run relationship of the selected variables, it is also interesting to know if a break occurs in this long-run relationship among the variables rather than in the individual series. For completeness, therefore, with the full sample (Jan. 1989-Jan. 2008), this study employs the most recent Johansen co-integration technique that allows for structural breaks at known points in time (Johansen et al. 2000). The results show that, with five lags (k=5), the trace tests reject the hypothesis of no co-integrating vector ($r = 0$), but fail to reject the null of one co-integrating vector ($r = 1$) at the 5% level (Table 2). This result suggests that there is one stable long-run equilibrium relationship among the four variables. In addition, the results of weak exogeneity test show that the commodity price, energy price and exchange rate are weakly exogenous at the 5% level (Table 3).

These findings indicate that these three variables are driving factors in the model and significantly affect the long-run movements of U.S. food prices, but are not influenced by U.S. food prices. Finally, using the relevant long-run coefficient (β_1), and normalizing the coefficient of the food price, the long-run equilibrium relationship among the variables is as follows:

$$FP_t = \underset{(1.16)}{0.02}\ EP_t + \underset{(10.66)}{0.18}\ CP_t - \underset{(4.18)}{0.12}\ ER_t + \underset{(3.08)}{0.002}\,trend + \underset{(16.6)}{0.02}\ E_{1,t} + \underset{(2.39)}{0.03}\ E_{2,t} \qquad (3)$$

[3] Since two co-integrating relationships are found with subsample II, an over-identification problem may arise because of the stationarity caused by the linear combination of the two co-integration relations (Harris and Sollis 2003). To solve this problem, we impose restrictions on the co-integrating spaces (β) to identify unique co-integrating vectors. We obtain the likelihood ratio statistic of 3.26 (p -value = 0.20), indicating that the over-identifying restrictions cannot be rejected.

Equation (3) shows that all variables except the energy price are statistically significant at the 5% level (parentheses are t-values) and have the expected signs. The results thus imply that over the 1989-2008 the commodity price and exchange rate have been significant factors influencing the food price. As compared with the finding from subsample II, this further confirms the strong linkage between energy and agricultural markets in recent years. Notice that the permanent dummy variables ($E_{1,t}$ and $E_{2,t}$) are found to be statistically significant at the 5% level in equation (3), implying that the dummy variables are play a significant role within the co-integrating relationship. In other words, this verifies that the structural break has a significant impact on the long-run relationship between the U.S. food price and its major determinants.

Analyzing Short-Run Relationships

To identify the short-run adjustment to long-run steady states, as well as the short-run dynamics between the food price and its major determinants, the VEC model is estimated with the identified co-integration relationships in subsample II.[4] The results of the parsimonious vector error-correction (PVEC) models show that the error-correction terms for food price and commodity price are negative and significant at the 5% significance level (Table 4). The negative coefficient of the error-correction term ensures that the long-run equilibrium can be achieved. The absolute value of the error-correction term indicates the speed of adjustment to equilibrium. The results indicate that, whenever deviating from equilibrium conditions, food price and commodity price adjust to correct long-run disequilibria in the U.S. agricultural market. However, the adjustment toward equilibrium is not instantaneous. For example, the commodity price adjusts by 34% to the first equilibrium (ec_1) and 5% to the second equilibrium (ec_2) in one month. These results imply that it takes approximately 3 months (1/0.34 = 2.94 months) and more than 20 months (1/0.05 = 20 months), respectively, to eliminate the disequilibria.

[4] The methodology used to find this representation follows a general-to-specific procedure (Hendry 1995). Specifically, since with subsample II the energy price and exchange rate are found to be weakly exogenous to the system, the VEC model is first estimated conditional on the two variables. By eliminating all the insignificant variables based on an F-test, the parsimonious VEC (PVEC) model is then estimated using full-information maximum likelihood (FIML) (Harris and Sollis 2003). The multivariate diagnostic tests on the estimated model as a system indicate no serious problems with serial correlation, heteroskedasticity, and normality (Table 4). Hence, the model specification does not violate any of the standard assumptions.

Table 4. Parsimonious VEC model with Subsample Π (Nov. 2001-Jan. 2008)

		ΔFP_t	ΔCP_t
Coefficient estimates	ΔFP_{t-1}	-0.67 (-6.01)**	
	ΔFP_{t-2}	-0.46 (-3.23)**	-1.04 (-1.49)
	ΔFP_{t-3}	-0.22 (-1.46)	1.61 (2.20)**
	ΔFP_{t-4}	-0.36 (-2.92)**	
	ΔFP_{t-5}	-0.19 (-1.60)	
	ΔCP_{t-1}	0.03 (-1.68)*	0.36 (3.41)**
	ΔCP_{t-3}	0.04 (2.02)**	0.10 (1.03)
	ΔCP_{t-5}		0.14 (1.70)*
	ΔER_t	-0.09 (-2.05)**	-0.34 (-1.70)*
	Constant	0.31 (2.71)**	0.59 (0.95)
	ec_{t-1}	-0.07 (-2.99)**	-0.34 (-2.90)**
	ec_{t-2}	-0.01 (-1.39)	-0.05 (-2.00)**
Diagnostic tests	Serial correlation	1.54 [0.11]	
	Heteroskedasticity	5.73 [0.22]	
	RESET	0.56 [0.99]	

Note: ** and * denote significance at the 5% and 10% levels, respectively. Parentheses are t-statistics. Brackets in diagnostic tests are p-values.

The coefficients of the lagged variables in the PVEC models show that the short-run dynamics (causal linkages) of the dependent variables. Specifically, the food price is negatively correlated with lagged food price and exchange rate, while it is positively correlated with commodity price. In addition, the commodity price is negatively correlated with exchange rate, but positively correlated with food price and lagged commodity price. Unlike the long-run results, however, energy price is found to have little effect on food and commodity prices in the short-run. Finally, the results show that there is a significant short-run dynamic effect between food price and commodity price; that is, the food price is significantly affected by lagged changes in commodity price, which is also influenced by the lagged changes in food price.

CONCLUDING REMARKS

While many studies have used descriptive methods to investigate factors that affect the recent surge in U.S. food prices, relatively little attention has been paid to the empirical analysis of this issue. In this paper, therefore, we attempt to examine the short- and long-run effects of changes in market factors such as prices of energy and agricultural commodities and exchange rate on changes in U.S. food prices. For this purpose, the Johansen co-integration analysis and VEC model are used to estimate monthly price data from 1989 to 2008. The results show that the agricultural commodity prices and exchange rate play key roles in affecting the short- and long-run behavior of U.S. food prices. We also find that while the energy price has been a significant factor influencing U.S. food prices in recent years in the long-run, it has little impact on U.S. food prices in the short-run. These findings further suggest that in the long-run, the linkage between energy and agricultural commodity/food markets has been recently quite strong, due mainly to crop-based bio-fuel production.

An important implication of our findings is that, since the exchange rate has a sizeable effect on the agricultural commodity price and thus food price, the uncertainty about commodity prices could increase as the value of the U.S. dollar fluctuates against currencies of other major trading partners. As the U.S. economy moves from recession to recovery, for example, appreciation of the U.S. dollar is likely to increase U.S. imports of agricultural goods through a decline in import prices, thereby resulting in a decrease in domestic commodity prices and hence food prices. Another important implication is that the Energy Security Act of 2005 and the Energy Independence of Security Act of 2007 has indeed resulted in record or near record prices for key commodities (e.g., corn, soybeans and wheat) through significant growth in the use of farm commodities for bio-fuel production. As such, as long as bio-fuel production continues under these legislations, agricultural commodity and food prices are expected to remain high in the future.

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APPENDIX 1
EMPIRICAL TIME-SERIES METHODOLOGY

The Johansen approach starts with an unrestricted vector autoregression (VAR) involving up to k lags of z_t:

$$z_t = \mu + A_1\Delta z_{t-1} + ... + A_k z_{t-k} + \varepsilon_t \tag{A1-1}$$

where z_t is a $(n \times 1)$ vector of endogenous variables; each of the A_k is an $(n \times n)$ matrix of parameters; μ is a vector of constant; and ε_t is white noise. This type of VAR model is well suited to estimate dynamic relationships among jointly endogenous variables and causal mechanisms without imposing strong a *priori* restrictions (Sims 1980, Harris and Sollis 2003). Because the right-hand side regressors are the same for all equations, ordinary least squares (OLS) is an efficient way to estimate equation (A1-1).

Following Johansen (1995), equation (A1-1) can be reformulated into a vector error-correction (VEC) form in order to impose the co-integration constraint as follows:

$$\Delta z_t = \mu + \Gamma_1\Delta z_{t-1} + ... + \Gamma_{k-1}\Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \tag{A1-2}$$

where $\Gamma_i = -(I - A_1 - ... - A_i)$ $(i = 1,...,k-1)$, which are the coefficient matrices of short-term dynamics; $\Pi = -(I - A_1 - ... - A_k)$ are the matrix of long-run coefficients; and Δ is the difference operator. Equation (A1-2) thus contains information on both the short-run ($\hat{\Gamma}_i$) and long-run ($\hat{\Pi}$) adjustments to change in z_t. The coefficient matrix Π can be decomposed into a matrix of weight, α, and a matrix of co-integration vector, β, that is $\Pi = \alpha\beta'$. The matrix α indicates the speed of adjustment to equilibrium and β' is a matrix of long-run coefficients such that the term $\beta' z_{t-k}$ represents up to $(n-1)$ co-integration relationships in the multivariate model, which ensures that the z_t converges with their long-run steady state solutions. Assuming that z_t is a vector of non-stationary $I(1)$ variables, then all the terms in equation (A1-2) that involve Δz_{t-i} are stationary, while Πz_{t-k} must be stationary for $\varepsilon_t \sim I(0)$ to be white noise, which thus the standard asymptotic results apply.

If all variables in z_t are co-integrated, equation (A1-2) can be reformulated as a short-run dynamic models to estimate the VEC model with the error-correction terms explicitly included as follows:

$$\Delta z_t = \mu + \Gamma_1\Delta z_{t-1} + ... + \Gamma_{k-1}\Delta z_{t-k+1} + \alpha(\beta' z_{t-1}) + u_t \tag{A1-3}$$

where $\beta' z_{t-1}$ is a measure of the error or deviation from the equilibrium, which is obtained from residuals from the co-integrating vectors. Equation (A1-3) thus incorporates both short-run and long-run effects. If the long-run equilibrium holds, for example, $\beta' z_{t-1} = 0$. During periods of disequilibrium, on the other hand, this term is non-zero and measures the distance of the system from equilibrium during time t; in this case, an estimate of α provides information on the speed-of-adjustment, which implies how the variable z_t changes in response to disequilibrium.

APPENDIX 2

TESTING FOR UNIT ROOTS UNDER THE STRUCTURAL SHIFTS

Perron (1989) develops a modified ADF test for the presence of a unit root with three alternative models to take into account structural changes in the deterministic trend function. Given a known structural break, the approach is generalized to allow one-time change in the structure occurring at a time T_B. Model A is referred to as the crash model and allows for a one-time change in the intercept of the trend function. Model B is known as the changing growth model and considers a change in the slope of the trend function without any sudden change in the intercept. Model C allows for both effects (slope and intercept) to take place simultaneously.

To validate or reject our graphical inspections (see Figures 5a and b), we follow Perron's (1989) models A and C by specifying the regression equations:

Model A: $y_t = \mu_0 + \mu_1 DU_t + \delta t + u_t$ (A2-1)

Model C: $y_t = \mu_0 + \mu_1 DU_t + \delta_0 t + \delta_1 DT_t + u_t$ (A2-2)

where $DU_t = 1$ if $t > T_B$ and 0 otherwise where T_B is the time at which the trend and/or intercept break occur(s); and $DT_t = t$ if $t > T_B$ and 0 otherwise. We estimate equations (A2-1) and (A2-2) using OLS regression for values of T_B in the neighborhood of visually inspected break dates. The results show that all regressions have relatively high R^2 values (0.71~0.99) (Appendix Table 1). The coefficients on the intercept, trend, and intercept- and trend shifts in the commodity price, energy price, exchange rate and ethanol production are statistically significant at the 5% level. In addition, the coefficients on the intercept, trend, and intercept shift in the food price are significant at the 5% level. The results thus indicate that the inclusion of the trend/intercept shifts is statistically important. For completeness, we use the estimated models to generate fitted values (solid lines) of the dependent variables. Figures 5a and b provide graphical validation of the structural changes obtained from the regression results.

Appendix Table 1. OLS regression results on structural shifts in selected variables

	Dependent variable				
Independent variable	$\ln FP_t$	$\ln CP_t$	$\ln EP_t$	$\ln ER_t$	$\ln ETH_t$
Intercept	4.315 (323.2)**	4.652 (501.0)**	4.354 (409.0)**	4.441 (769.0)**	4.099 (148.0)**
Intercept shift	-0.011 (-4.79)**	-0.145 (-10.8)**	-0.151 (-10.2)**	0.115 (12.4)**	-0.509 (-14.7)**
Trend	0.002 (124.0)**	0.001 (2.36)**	0.001 (7.71)**	0.001 (12.2)**	0.008 (14.1)**
Trend shift		0.003 (14.7)**	0.006 (24.3)**	-0.003 (-20.6)**	0.006 (10.3)**
Time of shift	Dec. 1998	Nov. 1998	Feb. 1998	Oct. 2001	Feb. 1996
n	229	229	229	229	229
R^2	0.99	0.71	0.95	0.75	0.96

Note: FP_t, CP_t, EP_t ER_t and ETH_t represent U.S. food price, U.S. commodity price, U.S. energy price, exchange rate, and U.S. ethanol production, respectively. ** denotes significance at the 5% level. t-values are given in parentheses. Both intercept and trend shifts occur in CP_t, EP_t ER_t and ETH_t, while only intercept shift occurs in FP_t.

Finally, in conducting unit root tests that allow a shift in the intercept (model A) and in the intercept and trend (model C), Perron (1989) estimates the following regressions:

$$\text{Model A: } y_t = \mu^A + \beta^A t + \theta^A DU_t + \gamma^A TB_t + \alpha^A y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + \varepsilon_t \quad \text{(A2-3)}$$

$$\text{Model C: } y_t = \mu^C + \beta^C t + \theta^C DU_t + \gamma^C TB_t + \delta^C DT_t + \alpha^C y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + \varepsilon_t \quad \text{(A2-4)}$$

where $TB_t = 1$ if $t = T_B$, and 0 otherwise; α is the autocorrelation between y_t and y_{t-1}; and summation ($\sum$) includes augmented difference terms, which is used primarily to generate cleaner estimation of the autocorrelation. The null of hypothesis of non-stationary implies that $\alpha = 1$. This can be tested by a t-test with critical values tabulated by Perron's Monte Carlo simulation (Perron 1989).