Market Integration for Shrimp and the Effect of Catastrophic Events

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Seasonal unit-root testing and seasonal cointegration methods are employed to investigate the price transmission in U.S. shrimp markets. ARIMA and Vector Error Correction Models (VECM) are used to identify the effect of catastrophic events on individual price series in one region and the spillover effects in the price series for other regions. Results showed that a cointegrating relation exists between neighboring states, specifically between Alabama and Mississippi and Louisiana and Texas. Cointegrating relations also exist between the Gulf States and the Pacific region, but not the Atlantic region, and the price of imported shrimp is cointegrated with each of the domestic shrimp price series. Finally, while Katrina had an effect on shrimp prices in Gulf States, the effect was not long lasting.

Key Words: catastrophic events, cointegration, market integration, seasonal unit-roots, spillover effects

JEL Classifications: C13, Q11, Q13

Introduction

The flow of goods and information across geographic regions may cause spatially separate markets to be integrated resulting in price transmission across regions. If a commodity price change in one region induces a directional price change in another region of equal degree, the markets in each region are said to be completely integrated as a single market. Spatial price integration in commodity markets at the national and international level has been extensively studied in the literature. Several papers have investigated the degree of market integration for spatially distinct agricultural markets through the use of cointegration tests. Examples for several agricultural commodities and markets include Ravallion (1986), Ardeni (1989), Goodwin and Shcroeder (1991), Alavalapati, Adamowicz and Luckert ((1997), Sarker (1993), Hänninen, Toppnen and Ruuska (1997), Toppinene and Tiovonen (1998), Murray and Wear (19998) and Prestemon and Holmes (2000). In these studies, the error correction mechanism that drives the cointegrating relationships is assumed to be the spatial arbitrage.

This paper investigates the price transmission in U.S. shrimp markets where markets are segmented by region. Cointegration tests are used to investigate whether different shrimp markets are integrated. Given potential international linkages, the price arbitrage between the domestic and import market is also considered. Prestemon and Holmes (2000) note that spatial arbitrage may not make sense for *in situ* resources. In their study of stumpage prices, they argue that cointegration for *in situ* resources can occur because of intertemporal (rather than spatial) arbitrage. We also employ the Prestemon and Holmes assumption that cointegration between the different shrimp markets occurs because of intertemporal arbitrage. Similarly, we assume that the different shrimp markets can be defined over the "information space" as those submarkets that respond statistically in a similar way to the same information about the factors affecting shrimp demand and supply.

In this study, U.S. shrimp markets are defined by the following regions: Atlantic, Pacific, and Gulf Coast Region. For the Gulf Coast, the market is further disaggregated by state (Louisiana, Mississippi, Texas, etc.). Given the importance of imports to U.S. shrimp supply, we also treat imports as a separate market. For the last decade, U.S. shrimp imports have increased nearly 11% per year on average and have gained a greater share of total U.S. supply. At the same time, shrimp prices saw a more than 75% decline. In 1996, domestic shrimp production and shrimp imports were 21% and 79% of total U.S. supply, respectively. Currently, imports account for about 90% of total U.S. shrimp supply (NOAA, 2007). Thus investigating the effect imports on domestic shrimp prices will provide a better understanding of the relationship between these markets.

In addition to investigating market integration, we also investigate the effect of catastrophic events on shrimp prices and whether these effects spill over from one submarket to another. In particular, we investigate the effect of hurricane Katrina on the shrimp prices in the Gulf Coast region and whether these effects are reflected in shrimp prices for the other regions. Although, shrimp is the leading seafood product in the United States where domestic production has faced significant international competition, few studies have analyzed the spatial price linkages in U.S. and international shrimp markets. A noted exception is Vinuya (2007) who used cointegration techniques to investigate if world shrimp prices share a common stochastic trend and if the law of one price held across international markets. His results show strong price linkages between the Japanese, American, and European shrimp markets with evidence supporting the law of one price.

The overall goal of this research is to gain a better understanding of the relationship between shrimp prices for the different markets specified in this study. Specific objectives are twofold. First, this work will investigate whether the different shrimp markets are integrated. Particular attention is given to the relationship between import and domestic prices. Second, this work will investigate the effect of catastrophic events on shrimp prices in one market and the spillover effect on shrimp prices in other markets.

Data and Empirical Procedure

The data used in this study consists of dockside shrimp prices for several species of shrimp for the Atlantic and Pacific regions and Gulf Coast states as well as imports. The Gulf Coast states include Alabama, Florida, (West coast only), Louisiana, Mississippi, and Texas. The frequency of observation is monthly and span from 1990 to 2008. Brown, pink and white shrimp compose over 99% of the shrimp harvest in the Gulf Coast states. Pink shrimp are not included in the analysis because of incomplete price series. Additionally, pink shrimp are more abundant during winter and spring. Thus, the price data for pink shrimp cannot be used to study the effect of hurricane Katrina, which occurred in the month of August. Data for brown shrimp can be used to study the effect of Katrina as brown shrimp are more abundant from June to October. However, brown shrimp price data for several states and regions are incomplete. Similarly, white shrimp are more abundant during late summer and fall months. In addition, the data are complete for all Gulf States and the Atlantic region with the exception of the Pacific region.

Therefore, we use the price data for white shrimp in our study. We also use price data for spot shrimp for the Pacific region, as this species has the largest percentage of harvest for this region. Eight series of shrimp prices are included in the analysis. Five are for shrimp from the Gulf Coast, one each for the Atlantic and Pacific regions and one for imports.

In previous work, the cointegration approach is used to test for and estimate the cointegrating relationship between the price series for agricultural products from different regions. The cointegration approach is applied to series that are non-stationary purely due to unit roots. In other words, the series are integrated of order one, denoted as I(1). The aim of the cointegration approach is to find linear combinations of variables that also remove unit roots. For the simple bivariate case, if y_t and x_t are both I(1), there may be a unique value of β such that

(1) $u_t = y_t - \beta_0 - \beta_1 x_t$

is I(0). In other words, there is no unit root in the linear combination of y_t and x_t. The term u_t is also referred to as an "error correction term" (Hamilton, 1994). The cointegration approach, therefore, assumes that the root of interest has a modulus that is precisely one and that it corresponds to a zero-frequency peak in the spectrum. Further, the cointegration approach assumes that there are no other unit roots in the system.

However, many economic time series exhibit substantial seasonality (Hylleberg et al. 1990). Therefore, there is a possibility that there may be unit roots at other frequencies, for example seasonal frequencies. Figure 1 shows the price series for white shrimp for Mississippi and Louisiana. It is clear that both series display strong seasonal patterns. Box and Jenkins (1970) implicitly assume that there are seasonal unit roots by using the seasonal differencing filter.

For annual data, the seasonal unit root model for the dependent variable ω_t is defined as:

$$(2)\quad \begin{aligned} \omega_t &= \varphi \kappa_t + \eta_t, \\ \eta_t &= \eta_{t-12} + \xi_t, \end{aligned}$$

where κ_t is an independent variable, η_t is the error term following a twelfth difference process, ξ_t is i.i.d. normal, φ is a parameter, and t represents the monthly observation. In this case, the seasonal differencing filter is applied as

$$(3)\quad \omega_t - \omega_{t-12} = \varphi(\kappa_t - \kappa_{t-12}) + \xi_t \ .$$

Seasonal unit roots have largely been used when forecasting was the primary focus (Clements and Hendry, 1997). One problem with the seasonal unit root model is that it is often rejected in empirical work (McDougall, 1995). Hylleberg et al. (1990) show that in the presence of seasonal unit roots, the standard procedure for testing for cointegration is inappropriate. The alternative strategy is to first test for seasonal unit roots. Then, appropriately filter the series according to the tests for seasonal roots. The seasonally adjusted series can then be used for estimation and testing for cointegration at zero frequency.

Hylleberg et al. (1990) develop a general procedure that can test for unit roots at some seasonal frequencies without maintaining that unit roots are present at all seasonal frequencies. Beaulieu and Miron (1993) extend the Hylleberg et al. procedure to monthly data.
We employ the Beaulieu and Miron (1993) extension of the Hylleberg et al. (1990) procedure for monthly data first to formally test the assumption of seasonal unit roots in the price series. Beaulieu and Miron (1993) suggest that to show that no seasonal unit root exists at any frequency, π_k in equation (4) must not equal zero for k=2 and for at least one member of each of the sets {3, 4}, {5, 6}, {7, 8}, {9, 10}, {11, 12}.

(4) $B(L)*W_{13t} = \sum_{k=1}^{12} \pi_k W_{k,t-1} + d_1 + \sum_{k=2}^{12} d_k Z_{kt} + \zeta_t$

where W_k are functions of current and lagged values of monthly shares (details are provided in the appendix), Z_k are dummy variables and $B(L)$ is a polynomial in the lag operator. Therefore, we perform the test using a two-sided t-test for k=2 and F-tests for the other sets. Equation (4) is estimated for the eight price series. Based on the results of seasonal unit roots tests, the price series are seasonally filtered to remove any existing seasonal roots.

Next we employ the Johansen (1992) method to test for the presence of cointegrating relationships between the different price series. The Johansen method consists of estimating a m-dimensional, k^{th} order VAR-model, written in error-correction form

(5) $\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-k} + \varepsilon_t$

where X is a vector of prices, and Γ_1 through Γ_{k-1} ($m \times m$) and Π ($m \times m$) are parameters to be estimated for some $r = 1, \ldots, m$. The errors ε are assumed to be independent and Gaussian with mean zero and covariance matrix Ω.

The rank of matrix Π is of interest with regard to the long-run cointegrating relationships between variables in the model. Engle and Granger (1987) provide a general definition of cointegration: if all the variables in X_t are integrated of order d, and there exist a cointegrating vector $\beta \neq 0$ such that $\beta' X_t$ is integrated of order d-r, then the processes in X_t are cointegrated of order CI(d,r). If the rank of Π is equal to m then all variables in X are stationary. If the rank of Π equals zero, model (5) reduces to a differenced vector time series model implying that no cointegration relationships exist among variables in X. If the rank of Π is greater than zero but less than m, there exist two matrices α ($m \times r$) and β ($m \times r$) such that $\Pi = \alpha\beta'$. β consists of r

cointegrating vectors representing the long-run relationship between the variables in X while the α's are the adjustment parameters following a deviation from the long-run relationships (Johansen and Juselius, 1990). Under the assumption of cointegration of order r model (5) can be written as

(6) $$\Delta X_t = \alpha\beta' X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \varepsilon_t$$

Equation (6) has the property that under suitable conditions on the parameters the process is non-stationary, ΔX_t is stationary, and $\beta' X_t$ is stationary, see Johansen (1992, 1995).
To determine the cointegrating rank Johansen (1992) proposes two possible tests: the λ_{max} test and the trace test[1]. The λ_{max} test is obtained using $\lambda_{max} = -T\, ln(1 - \lambda_{r+1})$ where the λ_i's are the eigenvalues of the matrix $\Pi = \alpha\beta'$. The idea behind the λ_{max} test is that if the $(r+1)^{th}$ eigenvalue is not different from zero then the smaller eigenvalues are also not different from zero. This is to test whether there exist r+1 cointegrating vectors against r cointegrating vectors.
The trace test is obtained using $\lambda_{trace} = -T \sum_{i=r+1}^{m} ln(1 - \lambda_i)$ where only the $m{-}r$ smallest non-zero eigenvalues, $\lambda = (\lambda_{r+1}, \ldots, \lambda_p)$, are used in the calculation of the test statistic. The null hypothesis for the trace test is that there are (m-r) cointegrating vectors. The trace test is used in this research since it provides a more consistent way of determining the cointegration order (Johansen, 1992, Johansen and Juselius, 1992).

We employ intervention analysis (Enders, pp. 240-247) to identify a change in the mean of a stationary time series. The impacts of a stochastic catastrophic event like Hurricane Katrina

[1] Cointegrated ADF (CADF) test of Engle and Granger (1987) is also an alternative test for cointegration.

can be measured *ex post* through the use of dummy variables. Let D_t^τ be a "pulse" dummy variable defined as

(7) $$D_t^\tau = \begin{cases} 0, & t \neq \tau \\ 1, & t = \tau \end{cases}$$

and let S_t^τ be a "step" dummy variable defined as

(8) $$S_t^\tau = \begin{cases} 0, & t < \tau \\ 1, & t \geq \tau \end{cases} .$$

The pulse function (7) is used to test for a short-run price impact associated with damage to the shrimp harvesting fleet and dock infrastructure due to Hurricane Katrina. The step function (8) is used to test for a structural adjustment due to long-term effects on the harvesting fleet and stock of shrimp.

We first investigate the effects of a catastrophic event in markets affected by the event, specifically, Louisiana and Mississippi. We use an ARMA(*L*,*Q*) process to model the dynamic properties of the price series

(9) $$p_t = f(p_{t-l}, \psi_{t-q}, D_t, D_{t-1}, \ldots, D_{t-4})$$

where *L* and *Q* represent respectively the order of the autoregressive and moving average process and ψ_t is a white noise error with mean zero and constant variance.

To conduct the intervention analysis we obtain the series of error-correction terms, v_t using (1) where the β parameters are estimated using (6). The series of error-correction terms contains information on the effects of the catastrophic event. As noted earlier, this series is stationary. An ARMA(L,Q) process is used to model the dynamic properties of this stationary series

(10) $v_t = f(v_{t-p}, \varepsilon_{t-q}, D_t, S_t) + \varepsilon_t$

where ε_t is a white noise error with mean zero and constant variance and L and Q represent respectively the order of the autoregressive and moving average process. We use an AR(1) specification of equation (10) to estimate the effects of a catastrophic event:

(11) $v_t = \phi_0 + \phi_1 v_{t-p} + \phi_2 D_t + \phi_3 S_t + \varepsilon_t$

where $0 < \phi_1 < 1$. Damage to the shrimp harvesting fleet suggests that ϕ_2 is negative, while damage to dock infrastructure suggests that ϕ_2 is positive. Katrina significantly damaged property and infrastructure important to fishing and aquaculture industries in Louisiana, Mississippi, and Alabama. The commercial shrimp industry was particularly affected where Katrina destroyed or severely damaged shrimp boats and shrimp processing and storage facilities throughout the region (Buck, 2005). The gradual recovery following the damage to the infrastructure suggests that ϕ_3 is positive. The effect of a catastrophic event on shrimp price is calculated as $(\phi_2 + \phi_3)$ during the supply pulse.

The dynamic effects of a catastrophic intervention can be obtained from the impulse response function (Enders)

(12) $v_t = \phi_0 /(1 - \phi_1) + \phi_2 \sum_{i=0}^{\infty} \phi_1^i D_{t-i} + \phi_3 \sum_{i=0}^{\infty} \phi_1^i S_{t-i} + \sum_{i=0}^{\infty} \phi_1^i \varepsilon_{t-i}$.

Equation (12) traces out the dynamic impacts of a catastrophic event on the time path of shrimp prices. The change in long-run equilibrium is $LR = \phi_3 /(1 - \phi_1)$.

Empirical Results

Seasonal Unit Root Tests

Results of the seasonal unit root testing are presented in table 1 which are the results of the hypotheses tests that π_2 (using a t-test) and at least one member of each of the sets $\{\pi_3, \pi_4\}$, $\{\pi_5, \pi_6\}$, $\{\pi_7, \pi_8\}$, $\{\pi_9, \pi_{10}\}$, and $\{\pi_{11}, \pi_{12}\}$ (using F-tests) are equal to zero. Given these results, we fail to reject the hypothesis that π_k equals zero for any frequency for all price series except for the $\{\pi_{11}, \pi_{12}\}$ set for the price series for all five Gulf states and the Atlantic region. A look at the individual t-tests (not reported here) for frequencies eleven and twelve indicates the presence of seasonal unit roots at frequency twelve for the price series for the five Gulf States and the Atlantic region. Therefore, a twelfth difference is applied to these price series to remove the seasonal unit root at that frequency.

Cointegration Tests

Results of the Johansen cointegration tests are reported in table 2. Cointegration tests are performed between each of the two affected states, Louisiana and Mississippi, and the remaining Gulf States and other regions. In addition, we report results of cointegration tests between imports and all other price series. Several observations are apparent from the results in table 2. First, cointegrating relations exist between neighboring states, specifically between Alabama and Mississippi and Louisiana and Texas. These results provide evidence in support of the hypothesis that spatial arbitrage occurs between neighboring states. Second, cointegrating relations exist between the Gulf States and the Pacific region, but not the Atlantic region. Thus, market integration also exists between the Gulf region and the Pacific region, but not between the Gulf region and the Atlantic region. Finally, the price of imported shrimp is cointegrated with each of the domestic shrimp price series. This important observation shows that all the different domestic markets for shrimp are integrated with the import market.

Intervention Analysis

Table 3 reports the results of the intervention analysis. Based on the cointegration results, the intervention analysis is conducted between the following markets, Louisiana and Texas, Pacific and Imports, and Mississippi and Alabama, Pacific and Imports. Results of table 3 show statistically significant short-run price drops for the pairs Louisiana and Texas and Mississippi and Alabama. Figure 2 displays the predicted path of the error correction terms for these two pairs of the markets. The price drop vanished for Alabama by the fifth month and in the sixth month there is a slight price increase. Starting in the seventh month, any long-run effect is eliminated. In the case of Texas, it takes up to nine months for the price decrease to vanish. After the ninth month the long-run effect is nonexistent. For the remaining pairs the short-run effect on price is positive and statistically significant. On the other hand, the long-run price effects are not statistically significant for any of the pairs of markets.

Table 4 reports the short-run and long-run effects on the price levels for the different market pairs. Price effect are reported in both the natural logs and in levels ($/lb). The short-run effects in levels ($/lb) are calculated using $p_{2005:9} * (1 - 1 / exp(\phi_2 + \phi_3))$ where, $p_{2005:9}$ is the actual price observed the month immediately after the hurricane Katrina, and ϕ_2 and ϕ_3 are parameter estimates reported in table 3. The long-run effects in levels ($/lb) are calculated using $(p_{2005:9} / exp(\phi_2 + \phi_3)) * (-1 + exp(\phi_3 /(1 - \phi_1)))$. The spillover effect from Louisiana to Texas in the short-run is a price decrease by $0.69 per pound (49 percent).

Similarly, the spillover from Mississippi to Alabama in the short-run is a decrease in price by $0.35 per pound (15 percent). The spillover effects for the other markets are positive and range from $0.11 per pound (8 percent) to $0.24 per pound (11 percent).

Conclusion

This paper investigates the price transmission in U.S. shrimp markets where markets are segmented by region and imports using cointegration techniques. In addition to investigating market integration, the effect of hurricane Katrina on shrimp prices was also examined. Results showed that a cointegrating relation exist between neighboring states, specifically between Alabama and Mississippi and Louisiana and Texas, cointegrating relations exist between the Gulf States and the Pacific region, but not the Atlantic region, and the price of imported shrimp is cointegrated with each of the domestic shrimp price series. Finally, while Katrina had an effect on shrimp prices in Gulf States, the effect was not long lasting.

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Table 1. Tests for Seasonal Unit Roots

State	π_2	$(\pi_3\ \pi_4)$	$(\pi_5\ \pi_6)$	$(\pi_7\ \pi_8)$	$(\pi_9\ \pi_{10})$	$(\pi_{11}\ \pi_{12})$
Alabama	9.26 [b]	3.27 [b]	5.72 [b]	9.78 [b]	9.45 [b]	1.49 [b]
	(0.0027) [c]	(0.0401) [c]	(0.0038) [c]	(0.0001) [c]	(0.0001) [c]	(0.2272) [c]
Florida	13.20 [a]	4.58 [b]	7.36 [b]	4.84 [b]	11.34 [b]	1.95 [b]
	(0.0004) [c]	(0.0113) [c]	(0.0008) [c]	(0.0101) [c]	(0.0001) [c]	(0.1454) [c]
Louisiana	6.45 [a]	4.53 [b]	3.72 [b]	7.14 [b]	5.20 [b]	0.57 [b]
	(0.0119) [c]	(0.0120) [c]	(0.0385) [c]	(0.0010) [c]	(0.0063) [c]	(0.5669) [c]
Mississippi	13.04 [a]	6.62 [b]	11.28 [b]	9.96 [b]	15.31 [b]	1.14 [b]
	(0.0004) [c]	(0.0016) [c]	(0.0001) [c]	(0.0001) [c]	(0.0001) [c]	(0.3204) [c]
Texas	3.98 [a]	9.42 [b]	3.31 [b]	4.50 [b]	6.99 [b]	3.56 [b]
	(0.0405) [c]	(0.0001) [c]	(0.0318) [c]	(0.0123) [c]	(0.0012) [c]	(0.0302) [c]
Atlantic	9.67 [a]	15.82 [b]	7.60 [b]	12.96 [b]	8.09 [b]	0.78 [b]
	(0.0022) [c]	(0.0001) [c]	(0.0007) [c]	(0.0001) [c]	(0.0004) [c]	(0.4596) [c]
Pacific	9.43 [a]	10.46 [b]	14.50 [b]	9.40 [b]	12.94 [b]	4.33 [b]
	(0.0024) [c]	(0.0001) [c]	(0.0001) [c]	(0.0001) [c]	(0.0001) [c]	(0.0145) [c]
Imports	10.12 [a]	13.17 [b]	11.88 [b]	5.21 [b]	17.69 [b]	7.28 [b]
	(0.0017) [c]	(0.0001) [c]	(0.0001) [c]	(0.0062) [c]	(0.0001) [c]	(0.0009) [c]

Note: [a] These are two-sided t-statistics.
[b] These are F-statistics.
[c] These are p-values.

Table 2. Cointegration Tests for Pairs of Shrimp Prices

Cointegration tests for the Catastrophic Event

Affected State	Other State/Region	Johansen Trace a	Johansen Trace b	Cointegrated?
Louisiana	Alabama	11.830	11.914	No
	Florida	39.305	39.529	No
	Texas	1.895	3.062	Yes
	Atlantic	36.280	37.973	No
	Pacific	0.895	5.031	Yes
Mississippi	Alabama	0.239	5.217	Yes
	Florida	53.933	54.050	No
	Texas	47.873	47.874	No
	Atlantic	14.372	14.430	No
	Pacific	0.746	5.129	Yes
	Cointegration tests Between Imports and Other Series			
Imports	Alabama	1.272	1.380	Yes
	Florida	1.965	2.029	Yes
	Louisiana	1.562	1.623	Yes
	Mississippi	1.810	1.886	Yes
	Texas	1.769	1.846	Yes
	Atlantic	1.509	1.612	Yes
	Pacific	1.929	3.418	Yes

Note: [a] These are Johansen's trace tests with an intercept in the cointegrating equation and no intercept in the VAR. The critical value for the five percent level of significance is 3.84.

[b] These are Johansen's trace tests with no intercept in the cointegrating equation and an intercept in the VAR. The critical value for the five percent level of significance is 9.13.

Table 3. Intervention Analysis Results for Shrimp Prices Following Hurricane Katrina

Affected State	Other State/Region	Intercept	v_{t-1}	D_t^{τ}	S_t^{τ}	Adjusted R^2
Louisiana	Texas	0.220***	-0.340***	-0.367**	-0.029	0.61
	Pacific	0.117**	-0.064**	0.089*	0.016	0.30
	Imports	0.105***	-0.086***	0.078**	0.003	0.26
Mississippi	Alabama	0.698***	-0.398***	-0.175*	0.034	0.54
	Pacific	0.531***	-0.562***	0.094*	0.017	0.38
	Imports	0.100***	-0.094***	0.075**	-0.013	0.28

Table 4. Short-Run and Long-Run Shrimp Price Effects Following Hurricane Katrina

Affected State	Other State/Region	Short-Run Effect (Natural Log)	Long-Run Effect (Natural Log)	Short-Run Effect ($/lb)	Long-Run Effect ($/lb)
Louisiana	Texas	-0.40**	-0.02	-0.69**	-0.05
	Pacific	0.11*	0.02	0.14*	0.02
	Imports	0.08**	0.00	0.11**	0.00
Mississippi	Alabama	-0.14*	0.02	-0.35*	0.06
	Pacific	0.11*	0.01	0.24*	0.02
	Imports	0.06**	-0.01	0.14**	-0.03

Figure 1. Shrimp Prices for Louisiana and Mississippi ($/lb).

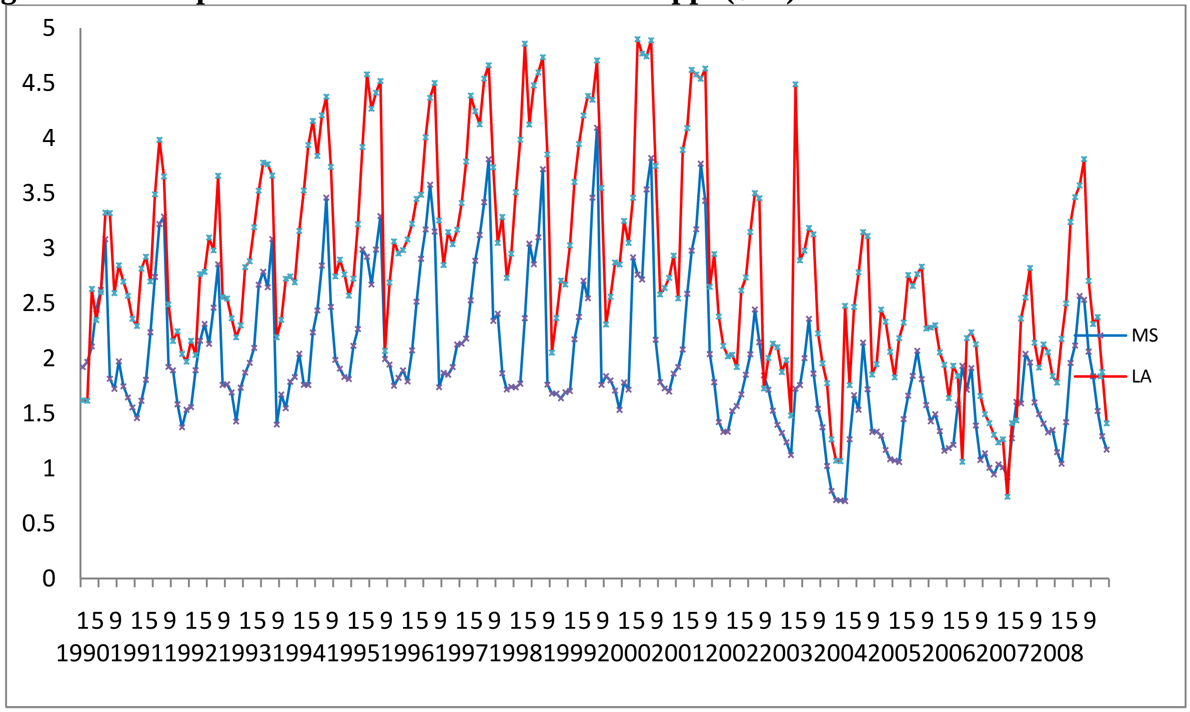
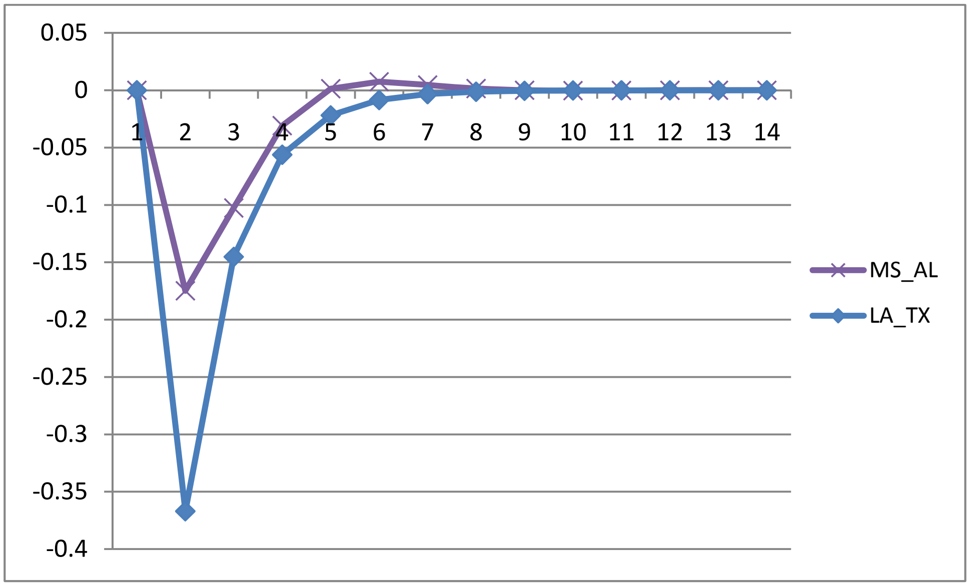


Figure 2. Impulse response Function Following Hurricane Katrina for Shrimp Prices in Alabama and Texas



Appendix

To test the assumption of seasonal unit roots we follow Beaulieu and Miron (1993) and assume that the first differences of monthly expenditure share, *m*, are represented by a general autoregression of the form:

(A1) $$B_m(L)w_m = \zeta_m, \quad m = 1,\ldots,M \ ,$$

where *B*(*L*) is a polynomial in the lag operator, *L*, and ζ is a white noise process. The testing procedure developed by Hylleberg et al. (1990) is based on linearizing the polynomial *B*(*L*) around the zero frequency unit root and the other *S*-1 seasonal unit roots. For monthly data, (*S*=12) the *S*-1 seasonal unit roots are:

(A2) $$-1;\ \pm i;\ -\frac{1}{2}(1 \pm \sqrt{3}\,i);\ \frac{1}{2}(1 \pm \sqrt{3}\,i);\ -\frac{1}{2}(\sqrt{3} \pm i);\ \frac{1}{2}(\sqrt{3} \pm i)$$

where $i = \sqrt{-1}$ and these roots corresponding to 6, 3, 9, 8, 4, 2, 10, 7, 5, 1, and 11 cycles per year, respectively (Beaulieu and Miron). Thus, linearization of *B*(*L*) (see Beaulieu and Miron for details) leads to:

(A3) $$B(L)^* W_{13t} = \sum_{k=1}^{12} \pi_k W_{k,t-1} + \zeta_t \ ,$$

where:

$$
\text{(A4)}\quad
\begin{aligned}
W_{1t} &= (1 + L + L^2 + L^3 + L^4 + L^5 + L^6 + L^7 + L^8 + L^9 + L^{10} + L^{11})w_t \\
W_{2t} &= -(1 - L + L^2 - L^3 + L^4 - L^5 + L^6 - L^7 + L^8 - L^9 + L^{10} - L^{11})w_t \\
W_{3t} &= -(L - L^3 + L^5 - L^7 + L^9 - L^{11})w_t \\
W_{4t} &= -(1 - L^2 + L^4 - L^6 + L^8 - L^{10})w_t \\
W_{5t} &= -\frac{1}{2}(1 + L - 2L^2 + L^3 + L^4 - 2L^5 + L^6 + L^7 - 2L^8 + L^9 + L^{10} - 2L^{11})w_t \\
W_{6t} &= \frac{\sqrt{3}}{2}(1 - L + L^3 - L^4 + L^6 - L^7 + L^9 - L^{10})w_t \\
W_{7t} &= \frac{1}{2}(1 - L - 2L^2 - L^3 + L^4 + 2L^5 + L^6 - L^7 - 2L^8 - L^9 + L^{10} + 2L^{11})w_t \\
W_{8t} &= -\frac{\sqrt{3}}{2}(1 + L - L^3 - L^4 + L^6 + L^7 - L^9 - L^{10})w_t \\
W_{9t} &= -\frac{1}{2}(\sqrt{3} - L + L^3 - \sqrt{3}L^4 + 2L^5 - \sqrt{3}L^6 + L^7 - L^9 + \sqrt{3}L^{10} - 2L^{11})w_t \\
W_{10t} &= \frac{1}{2}(1 - \sqrt{3}L + 2L^2 - \sqrt{3}L^3 + L^4 - L^6 + \sqrt{3}L^7 - 2L^8 + \sqrt{3}L^9 - L^{10})w_t \\
W_{11t} &= \frac{1}{2}(\sqrt{3} + L - L^3 - \sqrt{3}L^4 - 2L^5 - \sqrt{3}L^6 - L^7 + L^9 + \sqrt{3}L^{10} + 2L^{11})w_t \\
W_{12t} &= -\frac{1}{2}(1 + \sqrt{3}L + 2L^2 + \sqrt{3}L^3 + L^4 - L^6 - \sqrt{3}L^7 - 2L^8 - \sqrt{3}L^9 - L^{10})w_t \\
W_{13t} &= (1 - L^{12})w_t.
\end{aligned}
$$

In the presence of a constant and seasonal dummies, (A3) becomes:

$$
\text{(A5)}\qquad B(L)^* W_{13t} = \sum_{k=1}^{12} \pi_k W_{k,t-1} + d_1 + \sum_{k=2}^{12} d_k S_{kt} + \zeta_t \ .
$$

In order to test for seasonal unit roots, (A5) is estimated by ordinary least squares (OLS) and then the OLS test statistics are compared to the critical values provided Beaulieu and Miron. Testing for the various unit roots can be accomplished either through a one-sided or two-sided t-test or through an F-test. Beaulieu and Miron derive critical values for the t-statistics and the F-statistic. Based on their Monte Carlo results, they conclude that the F-statistics are better than the t-statistics. Beaulieu and Miron suggest that to show that no seasonal unit root exists at any frequency, π_k must not equal zero for k=2 and for at least one member of each of the sets {3, 4}, {5, 6}, {7, 8}, {9, 10}, {11, 12}. Therefore, we perform the test using a two-sided t-test for k=2 and F-tests for the other sets.