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Impact of Mandatory Price Reporting on Hog Market Integration

We examine whether mandatory price reporting (MPR), which is intended to facilitate transparent pricing, has impacted pricing relationships among U.S. hog markets. Hog markets are cointegrated both prior to and following enactment of MPR, but are not fully integrated in either period. That is, prices at alternative locations do not adjust one-for-one with price changes in other locations. Further, markets adjust to price shocks in other locations more slowly following MPR, which may be a coincidence associated with decreases in the proportion of spot market hog transactions.

Keywords: hog markets, cointegration, mandatory price reporting, market integration, regime shift.

Introduction

The U.S. hog industry, like other livestock/poultry industries, has experienced substantial consolidation and growth in alternative marketing arrangements since the early 1990s when spot transactions dominated trade (Grimes and Plain 2005, 2007). With lower quantities of livestock traded in spot markets, voluntarily reported prices made publically available by the USDA's Agricultural Marketing Service (AMS) are increasingly scrutinized as being unreliable or unrepresentative of industry trade. Congress passed the Livestock Mandatory Reporting Act of 1999 to facilitate transparent price discovery and encourage competition (Azzam 2003; Pendell and Schroeder 2006). Prior research by Fausti and Diersen (2004) on cattle suggests that voluntary price reporting contributes as much to price transparency as mandatory systems, but Pendell and Schroeder (2006) find that cattle markets became more fully integrated following enactment of MPR. Mandatory hog price reporting (MPR) went into effect in April 2001.

The objective of this study is to examine how MPR has influenced spatial market integration among four domestic spot markets for hogs (Iowa-Southern Minnesota; Peoria, Illinois; St. Joseph, Missouri; and St. Paul, Minnesota) and also examines the degree of integration with futures prices. With more complete price and transaction data available, integration between spatially dispersed markets may be expected to increase.

There exists a rich body of literature on spatial market integration in agriculture with empirical applications extending from crops (e.g., Goodwin 1992a; Brester and Goodwin 1993; Goodwin and Piggott 2001; Franken, Parcell, Sykuta, and Fulcher 2005) to cattle (e.g., Schroeder and Goodwin 1990; Pendell and Schroeder 2006) to hogs (e.g., Benson, Faminow, Marquis, and Sauer, 1994; Schroeder and Goodwin 1991). While studies examine the impact of MPR on spatial cattle price relationships (Fausti and Diersen 2004; Pendell and Schroeder 2006) and lamb price risk (Marsh and McDonnell 2006), no such research on hog markets exists.

Research on integration between spot and futures hog markets is mixed with earlier studies finding a lack of price integration (i.e., Schroeder and Goodwin 1991) and more recent studies finding stronger price integration relationships (i.e., Yang, Bessler, and Leatham 2001; Carter and Mohapatra 2008). In general, these studies suggest that futures lead spot prices, which is consistent with market efficiency theory. These relationships are revisited in the current study, and impacts of MPR on spatial price relationships are investigated.

Three daily U.S. spot price series and nearby hog futures prices for 1992 through August 2009 are analyzed. Following Pendell and Schroeder (2006), bivariate and multivariate cointegration tests credited respectively to Engle and Granger (1987) and Johansen (1988) are used to investigate long-run price relationships, and Gregory and Hansen's (1996) bivariate cointegration test, which allows for a regime shift, is used to ascertain influences of MPR on price integration. Then, Vector Error Correction speed-of-adjustment coefficient before and after MPR are examined. As no recent research examines spatial price cointegration in hog markets, this study fills a gap in the literature by providing insights regarding the impacts of MPR.

Previous Research

A considerable body of research has been conducted on market integration issues for numerous commodities both domestically and internationally (e.g., Goodwin 1992a; Goodwin 1992b; Goodwin and Piggott 2001; Franken, Parcell, Sykuta, and Fulcher 2005; Schroeder and Goodwin 1990; Pendell and Schroeder 2006; Benson, Faminow, Marquis, and Sauer, 1994).[1] This section reviews selected studies and emphasizes research on livestock industries with mandatory price reporting (MPR).

Several studies of spatial price relationships, including the most comprehensive investigation of the effects of MPR to date (i.e., Pendell and Schroeder 2006), have been conducted on cattle markets. Weekly fed cattle price relationship during the 1970s and 1980s were examined for four regions by Bailey and Brorsen (1985) in a multivariate autoregressive framework, for eight direct and terminal markets by Koontz, Garcia, and Hudson (1990) using Granger causality analysis, and for eleven direct and terminal markets by Schroeder and Goodwin (1990) using vector autoregression (VAR) models. Bailey and Brorsen (1985) found that Texas Panhandle led prices in other regions with feedback from the Omaha market. Koontz, Garcia, and Hudson (1990) found direct markets to dominate with the Nebraska direct market being the most influential. Schroeder and Goodwin (1990) also found larger volume markets in major cattle feeding regions to be dominant price discovery locations. Further, prices at larger volume markets fully adjusted to changes in other markets within one to two weeks, while smaller markets took two to three weeks.

Goodwin and Schroeder (1991) found limited cointegration, especially among geographically dispersed regional fed cattle markets, and found cointegration increasing over time with increased concentration in cattle slaughtering in the 1980s, which may have reflected decreasing trade and information costs or collusive/coordinated, noncompetitive basing-point pricing.[2] Goodwin's (1992b) VAR models and impulse response functions confirmed gradual structural change in U.S. cattle markets from the mid 1970s through early 1980s with subsequently faster price adjustment. In a VAR analysis of early 1990s dressed fed cattle prices, Schroeder (1997) found a faster speed of adjustment in fed cattle prices associated with processing plants in close proximity and a decrease in speed of adjustment for larger plants and plants with fewer cash transactions.

Fausti and Diersen (2004) examined the relationship between daily fed cattle prices reported for Nebraska direct voluntarily and South Dakota mandatorily from September 1999 through March 2001, and concluded that voluntary price reporting was as transparent for price discovery as the mandatory system. In contrast, Pendell and Schroeder (2006) found using standard and regime shift cointegration models that spatial market integration among five U.S. regional markets from January 1992 through June 2006 increased with enactment of MPR in April 2001. Ward (2008) examined relationships among spot prices and prices under alternative marketing arrangements (AMAs) for both cattle and hogs since MPR, and found using regression analysis that cash prices lead AMAs in rising markets and trail them in declining markets. Building on Ward's (2008) analysis of beef and hog price relationships, Lee, Ward, and Brorsen (2010) found that cash and AMAs prices are cointegrated with all but one procurement method. While bidirectional causality was found between some procurement prices, cash prices Granger cause all other procurement prices, indicating that cash markets remain of central importance to price discovery.[3]

In general, few hog price integration studies have examined spatial relationships (e.g., Faminow and Benson 1990; Benson, Faminow, Marquis, and Saur 1994; Chen and Lee 2008), as the focus has been on relationships between cash and futures prices (e.g., Schroeder and Goodwin 1991; Yang, Bessler, and Leatham 2001; Liu 2005; Carter and Mohapatra 2008), and as discussed above, cash and AMA prices (e.g., Ward 2008; Lee, Ward, and Brorsen 2010). Moreover, none of these studies examine hog price integration among multiple U.S. markets.

Schroder and Goodwin (1991) investigated short- and long-run price relationships between Omaha cash and CME futures daily prices for live hogs. They found that price discovery generally originated in the futures market with little short-run feedback from Omaha to futures. Despite the causality evidenced by short-run dynamics, the two markets operated somewhat independently, as long-term basis was generally nonstationary, suggesting that the cash price reflected futures prices in addition to variations in short-term demand and supply while futures responded to information expected to impact subsequent cash prices. The authors also noted that divergence between cash and futures markets was larger for contracts further from maturity. Yang, Bessler, and Leatham (2001) also found cointegration among cash and futures prices for hogs. Whereas futures prices provided unbiased estimates for most storable commodities, futures prices were biased estimates of cash prices for live and feeder cattle, and to a lesser extent hogs, which typically involve more regular storage patterns of processed products than beef. Carter and Mohapatra (2008) also detected cointegration of cash and futures hog prices from 1998 to 2004 and demonstrated that the CME futures is the primary price discovery point and that futures are unbiased predictors of cash prices except for distant contracts.

Empirical Methods and Procedures

Bivariate and multivariate time-series procedures are employed to examine price linkages and price responsiveness among spatially dispersed cash markets and the futures market for hogs (i.e., cointegrated price series will not diverge from one another in the long-run). The methods demonstrated here follow from Pendell and Schroeder (2008). The analysis begins with Augmented Dickey-Fuller (ADF) tests to determine whether individual price series are nonstationary (i.e., a unit root exists). If the null hypothesis of a unit root is not rejected for the data in levels (i.e., nonstationarity) but is rejected for the data in first differences (i.e. stationarity), then long-run equilibrium relationships may be estimated.[4] The well-known test for cointegration between two spatial markets attributed to Engle and Granger (1987) is estimated by ordinary least squares (OLS) as:

(1) *Model I, Standard Cointegration*: $Y_t = \alpha_0 + \alpha_1 Z_t + e_t$,

where Y_t and Z_t are individual price series, α_0 and α_1 are intercept and slope coefficients and e_t is the error term. If an ADF test for stationarity of e_t indicates the presence of a unit root (i.e., e_t is nonstationary), then the two price series are not cointegrated.

Multivariate tests of cointegration commonly employ the Johansen (1988) method, which utilizes trace and maximum eigenvalue tests to investigate the number of cointegration vectors (Enders 1995).[5] 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 meaning that all prices are pairwise cointegrated; but if more than one common trend exists, the price series are not fully integrated. Correspondingly, the null hypothesis for both tests is that there are no more than r cointegrating vectors. The alternative hypothesis for the trace test statistic is that there exist more than r cointegration vectors. The alternative hypothesis for the maximum eigenvalue test statistic is that there are exactly $r + 1$ cointegration vectors. To account for the possibility that MPR caused a structural change in long-run price relationships, a set of residual-based cointegration tests, developed by Gregory and Hansen (1996) to allow for potential regime shifts, are estimated using OLS as follows:

(2) *Model II, Regime Shift Cointegration*: $Y_t = \alpha_0 + \alpha_1 D_t + \alpha_2 Z_t + \alpha_3 Z_t D_t + e_t$,

where Y_t, Z_t, and e_t are defined as above; D_t is a binary dummy defined as 1 following MPR and 0 prior to MPR; α_0 is the intercept prior to MPR and α_1 represents the change in the intercept after implementation of MPR; α_3 is the slope coefficient prior to MPR and α_3 denotes the change in slope after implementation of MPR. As in *Model I*, an ADF test for stationarity of e_t from *Model II* is used to test for cointegration. However, standard ADF critical values are not appropriate for *Model II*, and the appropriate critical values are reported in Gregory and Hansen (1996).

Estimating equations (1) and (2) enables testing of several hypotheses. First, if both specifications indicate that all prices are (or are not) cointegrated, then MPR did not notably affect long-run equilibrium relationships among the markets. Second, coefficient estimates allow comparison of pre- and post-MPR levels of market price integration. For instance, if α_3 is (is not) statistically different from zero, then price relationships changed (did not change) with MPR. Furthermore, comparing estimates of α_2 to $\alpha_2 + \alpha_3$ reveals whether prices move more or less on a one-for-one (i.e., perfectly integrated) basis after MPR relative to before.

Because of the multitude of supply-demand factors impacting the hog market, equations (1) and (2) were estimated as a special case of the Vector Autoregressive (VAR) specification allowing for cross-correlation of the errors and simultaneity between price integration equations.

(3) $\Delta P_t = \alpha_0 + \alpha_1 D_t + \alpha_2 Z_t + \alpha_3 Z_t D_t + \sum \beta_{11}(k)\, \Delta P_{t-k} + \sum \beta_{12}(k)\, \Delta Z_{t-k} + \mathbf{\Omega}_{\mathrm{it}}$,

where t refers to time ($t = 1, 2, \ldots, T$), which for this study is weeks; k is the number of lag lengths; and $\mathbf{\Omega}$ is an n×1 vector of normally distributed random errors. The specification of (3) allows for efficient standard errors and unbiased coefficients in the hypothesis tests of α_2 and $\alpha_2 + \alpha_3$ statistically different from one, while accounting for simultaneity between price locations.

To further analyze the level of price relationship between spot market locations, error correction VAR,or Vector Error Correction (VEC), models, incorporating the binary MPR dummy D described above, are estimated to investigate whether the speed of price responsiveness among locations differs before and after MPR. Highly integrated markets quickly return to long-run equilibrium following price shocks (Enders 1995). The VEC model is specified as:

(4) $\Delta P_t = \beta_0 + \beta_1 \hat{e}_{t-1} + \beta_2(\hat{e}_{t-1} \times D) + \sum \beta_{11}(k)\, \Delta P_{t-k} + \sum \beta_{12}(k)\, \Delta Z_{t-k} + \lambda_t$,

where variables and subscripts are as defined in equation (3), and λ is a n×1 vector of normally distributed random errors. The first two terms following the intercept on the right-hand side of equation (4) are the speed-of-adjustment measure and an interaction term between the speed-of-adjustment measure and the binary MPR dummy variable. The lagged error terms specified in equation (4) are obtained from the OLS estimation of equation (1). The next two terms are lagged price variables following from the standard VEC model. A speed-of-adjustment coefficient close to one in absolute value indicates quick adjustment to deviations from equilibrium, whereas a value near zero indicates slow adjustment. If MPR improves availability of reliable price information, then speed-of-adjustment coefficients nearer to one in absolute value should be observed post-MPR.

Data

Data analyzed were for four weekly U.S. spot price series and nearby hog futures prices for 1992 through August 2009 are analyzed. U.S. spot prices for terminal markets in St. Joseph, Missouri; Peoria, Illinois; and St. Paul, Minnesota are obtained from Plain (2010). The Iowa-Southern Minnesota interior market prices are obtained from the Livestock Market Information Center (LMIC). Chicago Mercantile Exchange (CME) lean hog futures prices are obtained from the Commodity Research Bureau (CRB). Beginning with the February 1997 contract, the CME replaced its Live Hog Futures Contract with a Lean Hog Futures contract priced on a carcass basis (Wellman 1996). With implementation of MPR in April 2001, Iowa-Southern Minnesota and other markets began reporting prices on a carcass basis. LMIC adjusted pre-MPR live hog prices for Iowa-Southern Minnesota to reflect lean value, and this adjustment was also applied to the other price series investigated here.[6] The lean value adjusted prices are graphed in Figure 1 and appear to trend fairly closely together.

Summary statistics are reported for the hog prices on a carcass basis (Table 1). The correlations among prices are all above 0.90 with the exception of Peoria's correlations with CME and Iowa-Southern Minnesota, which is 0.89 in both cases.

Results

Prior to market integration analysis, Dickey-Fuller (DF) tests of stationarity were performed. The appropriate lag structure for the DF tests and all subsequent models was determined by minimizing the Akaike Information Criteria (AIC), and the lag length was set to four. In all cases, the null hypothesis of nonstationarity could not be rejected at the five percent confidence level. Thus, the price series were deemed nonstationary. First-differencing the data corrected for nonstationarity, meaning that the time series are integrated of order 1, denoted I(1).

Pre- and post-MPR Johansen unrestricted cointegration rank test statistics (Enders 1995) are reported in Table 2. Trace statistics computed from characteristic roots (i.e., eigenvalues) reject the null hypothesis of no cointegrating vector at the five percent level for each case. Hence, each market pair is deemed cointegrated prior to and following the enactment of MPR, meaning that long-run price relationships exist between these markets in both time periods.

Following Pendell and Schroeder (2006), adapted VAR models were estimated following equation (3) with a post-MPR dummy (= 1 post-MPR, = 0 o.w.) and this dummy interacted with prices included as exogenous variables (Table 3). If MPR enhances the availability of reliable price information, then markets may adjust *more fully* to price shocks in other locations after its enactment. However, the results do not indicate *full integration* (i.e., a one-for-one relationship) among these markets either before or after the enactment of MPR (Table 3). Specifically, the null hypothesis that the price coefficient equals one ($\alpha_2 = 1$) is rejected in most cases, as is the null hypothesis that the sum of the price coefficient and the coefficient on the price×MPR dummy interaction equals one ($\alpha_2 + \alpha_3 = 1$).[7] This finding is largely consistent across ordering of dependent and independent variables, and hence, we present the results for only one ordering of each market pair to conserve space. Unreported results are available from authors upon request.

The presence of a cointegrating relationship (Table 2) justifies an error correction VAR (or VEC), as opposed to the standard VAR model (Enders 1995). Speed-of-adjustment coefficients from three VEC models are reported in Table 4. Here, we focus on relationships among cash price series, for which MPR should be most important, as opposed to relationship with futures prices. VEC 1 and VEC 2 are standard VEC models applied to the entire sample and the pre-MPR portion, respectively. As such, VEC 1 does not account for the impacts of MPR. Here, speed-of-adjustment coefficients are statistically significant at conventional levels only for pairings of St. Joseph and St. Paul. For St. Paul/St. Joseph, in particular, the absolute value of the speed-of-adjustment coefficient -0.2295 indicates that about 23 percent of the price adjustment in St. Joseph is realized in St. Paul within one week. VEC 2 pertains to the same price relationships for the period prior to MPR, and in general suggests faster and statistically more significant adjustment to price shocks in that period. VEC 3 examines the impact of MPR more closely by interacting a binary MPR dummy with the lagged error terms as in equation (4). Here, column A indicates that 48 percent of the price adjustment in St. Joseph is realized in St. Paul within one week, whereas column B indicates that after MPR only 29 percent of the adjustment is realized within one week. In general, slower adjustment if observed after MPR, which seems counterintuitive if MPR increases information availability. Perhaps, other factors, such as the decreasing percentage of hog sales in negotiated (i.e., spot) markets, are contributing to such findings.

Figure 2 illustrates a positive relationship between the declining percentage of hog sales made in negotiated market transactions and annual maximum eigenvalue cointegration test statistics. The correlation between these two series is 0.3284. Following Godwin and Schroeder (1991) and Brester and Goodwin (1993), who regressed annual cointegration test statistics on market concentration measures, a bootstrapped regression of the test statistics on the percentage of negotiated sales is performed. However, no statistically significant relationship is detected.

Conclusions

This paper investigates price relationships among various U.S. hog markets and whether mandatory price reporting (MPR), which is intended to facilitate transparent price discovery, has had any detectable impact on these price relationships. Previous research on the impacts of MPR in cattle markets is mixed. Fausti and Diersen (2004) conclude that voluntary price reporting is as transparent for price discovery as the mandatory system, while Pendell and Schroeder (2006) find U.S. regional cattle markets to be more fully integrated after enactment of MPR.

As in Pendell and Schroeder's (2006) analysis of cattle markets, we also find that hog markets are highly cointegrated both prior to and following the enactment of MPR. Whereas Pendell and Schroeder find that cattle markets become more fully integrated after enactment of MPR, hog markets are integrated but not fully integrated in pre- and post-MPR periods. The unsurprising exception is that St. Paul, Minnesota terminal price is fully integrated with the Iowa-Minnesota regional price in both periods, meaning that the former responds one-for-one to changes in the later. Lack of full integration may reflect transaction costs and require threshold analysis beyond the current scope of this research (Goodwin and Schroeder 1991; Goodwin and Piggott 2001).

Further, vector error correction models indicate that hog markets appear to be responding to shocks in other locations more slowly post-MPR. This counterintuitive result may reflect a number of other events coinciding with enactment of MPR. For instance, the proportion of hog sales transacted through spot market exchanges has declined substantially in the last two decades, and correspondingly, may have adversely affected market responsiveness to price changes in other locations. It may be that MPR offsets such potentially adverse effects, but difficulty in disentangling such coinciding effects inhibits our ability to discern this possibility.

Table 1. Summary Statistics and Correlations for Futures and Spot Market Hog Prices

	Summary Statistics				Correlations				
	Mean	Max	Min	Std. Dev.	CME	IA-S.MN	St. Paul	St. Joseph	Peoria
CME	61.47	87.68	23.03	10.44	1.00				
IA-S.MN	60.47	90.43	14.19	11.66	0.93	1.00			
St. Paul	56.62	84.73	15.03	11.17	0.93	0.94	1.00		
St. Joseph	56.00	83.92	13.38	11.16	0.92	0.94	0.99	1.00	
Peoria	55.32	84.32	13.92	11.06	0.89	0.89	0.98	0.98	1.00

Note: n = 922 observations.

Table 2. Summary of Johansen Unrestricted Cointegration Rank Test Statistics

Dependent/Independent	Before MPR		After MPR	
	Eigenvalue	Trace Statistic	Eigenvalue	Trace Statistic
IAMN/Peoria	0.016**	17.118	0.024**	18.637
IAMN/St. Joe	0.016**	19.104	0.040**	26.038
IAMN/St. Paul	0.016**	18.934	0.035**	23.698
IAMN/Futures	0.054**	65.731	0.091**	50.275
St. Joe/Peoria	0.018**	30.489	0.032**	24.355
St. Joe/St Paul	0.071**	82.618	0.071**	41.644
St. Joe/Futures	0.048**	60.119	0.063**	37.249
St. Paul/Peoria	0.018**	29.503	0.037**	25.980
St. Paul/Futures	0.045**	58.596	0.054**	33.717
Peoria/ Futures	0.026**	38.911	0.037**	26.040

Note: N = 483 and 439 for the samples before and after MPR, respectively. ** indicates statistical significance at the 5% level. Lag length is set to 4. Trace test statistic critical value is 15.495.

Table 3. VAR Parameter Estimates from Regime Shift Model (Model II)

Dependent Market/ Independent Market	Constant	α_2 State	α_1 Post-MPR Dummy	α_3 Post-MPR Regime	H_0: $\alpha_2 = 1$ (p-Value)	H_0: $\alpha_2 + \alpha_3 = 1$ (p-Value)
IA-S.MN/CME	-0.2036	0.4558***	-1.1425	0.0309	0.0000	0.0000
	(0.8035)	(0.0268)	(1.1954)	(0.0209)		
St. Joseph/CME	-0.0215	0.4472***	-0.7918	0.0263	0.0000	0.0000
	(0.8391)	(0.0280)	(1.2483)	(0.0218)		
St. Paul/CME	0.4147	0.5056***	-1.2056	0.0246	0.0000	0.0000
	(0.8762)	(0.0292)	(1.3035)	(0.0228)		
Peoria/CME	-0.0898	0.4436***	-0.6486	0.0023	0.0000	0.0000
	(0.7856)	(0.0262)	(1.1688)	(0.0204)		
St. Joseph/IA-S.MN	0.6161	0.8950***	-0.3425	0.0108	9.7434×10^{-10}	3.0520×10^{-08}
	(0.4521)	(0.0172)	(0.6299)	(0.0108)		
St. Paul/IA-S.MN	0.9077**	0.9920***	-0.1668	-0.0052	0.6215	0.4105
	(0.4270)	(0.0162)	(0.5949)	(0.0102)		
Peoria/IA-S.MN	0.3838	0.8656***	0.0473	-0.0196**	0.0000	0.0000
	(0.4176)	(0.0159)	(0.5819)	(0.0100)		
St. Joseph/St. Paul	-0.2611	0.8451***	0.2585	0.0050	0.0000	0.0000
	(0.4094)	(0.0143)	(0.5441)	(0.0099)		
Peoria/St. Paul	-0.4201	0.8237***	0.4106	-0.0229***	0.0000	0.0000
	(0.3596)	(0.0125)	(0.4779)	(0.0087)		
St. Joseph/Peoria	0.1427	0.9429***	0.2115	0.0174*	0.0004	0.0158
	(0.4187)	(0.0161)	(0.5561)	(0.0103)		

Note: N = 922. ***, **, * denote statistical significance at 1%, 5%, 10% levels, respectively. Standard errors are in parentheses. Lag length is set to 4.

Table 4. Speed-of-Adjustment Coefficients from Vector Error Correction Models

	VEC 1	VEC 2	VEC 3		
	Speed-of-adjustment coefficient (entire period)	Speed-of-adjustment coefficient (pre-MPR)	Speed-of-adjustment coefficient (entire period) (A)	Size of speed-of-adjustment after MPR (B)	Net impact (= A + B)
St. Joseph/IA-S.MN	-0.0347	-0.1292**	0.0536	-0.1698**	-0.1162
	(0.0235)	(0.0565)	(0.0499)	(0.0847)	
St. Paul/IA-S.MN	-0.0363	-0.1732***	0.0218	-0.1076	-0.0858
	(0.0252)	(0.0637)	(0.0552)	(0.0911)	
Peoria/IA-S.MN	-0.0215	-0.1152**	0.0202	-0.0750	-0.0549
	(0.0166)	(0.0468)	(0.0381)	(0.0618)	
IA-S.MN/St. Joseph	-0.0159	-0.0292	-0.0532	0.0643	0.0112
	(0.0218)	(0.0790)	(0.0563)	(0.0897)	
IA-S.MN/St. Paul	-0.0131	-0.0036	-0.0257	0.0216	-0.0041
	(0.0220)	(0.0834)	(0.0574)	(0.0911)	
IA-S.MN/Peoria	-0.0172	-0.1242*	-0.0381	0.0338	-0.0043
	(0.0158)	(0.0715)	(0.0441)	(0.0665)	
St. Joseph/St. Paul	-0.1771*	-0.0955	-0.0629	-0.1411	-0.2041
	(0.0963)	(0.1868)	(0.1694)	(0.1722)	
St. Paul/St. Joseph	-0.2295**	-0.4702**	-0.4774**	0.2946	-0.1828
	(0.1031)	(0.2129)	(0.1901)	(0.1900)	
St. Joseph/Peoria	-0.0689	-0.3997**	-0.1088	0.0529	-0.0559
	(0.0462)	(0.1824)	(0.1155)	(0.1402)	
Peoria/St. Joseph	-0.0402	-0.4187*	-0.0408	0.0008	-0.0400
	(0.0441)	(0.2192)	(0.1122)	(0.1378)	
St. Paul/Peoria	-0.0669	-0.4687***	-0.2058*	0.1978	-0.0080
	(0.0497)	(0.1738)	(0.1137)	(0.1456)	
Peoria/St. Paul	-0.0418	-0.1378	0.084901	-0.18702	-0.1021
	(0.0451)	(0.1604)	-0.1004	-0.13242	

Note: N = 922. ***, **, * denote statistical significance at 1%, 5%, 10% levels, respectively.
Standard errors are in parentheses. Lag length is set to 4.

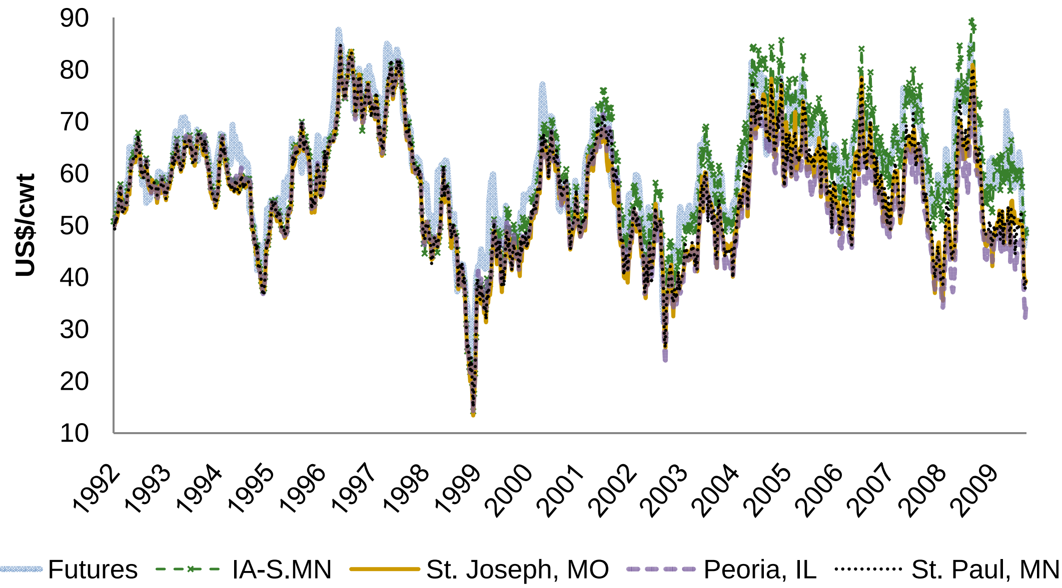
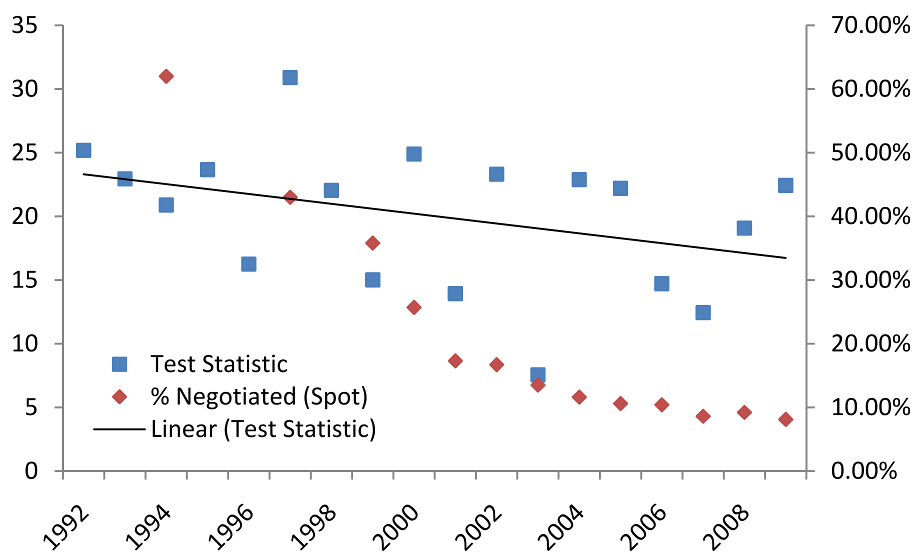


Figure 1. U.S. Hog Carcass Prices, 1992 -2009



Note: Bootstrapped regression: Test Statistic = 16.6390 + 12.4451×Percent Spot + error
(3.5924) (25.4523)

Figure 2. Maximum Eigenvalue Cointegration Test Statistics and Percentage of Hog Sales Negotiated in Spot Markets.

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Endnotes

[1] Markets perform efficiently when they are integrated (i.e., when the price in the receiving market equals the price in the shipping market plus transportation and other transfer costs associated with the trade).

[2] In a long-term equilibrium, prices are cointegrated. Highly cointegrated markets indicate strong spatial linkages.

[3] This finding does not really address the main question of whether thinning cash markets still offer base prices that are representative of supply and demand for quality animals.

[4] Cointegration necessitates that each of the time series be integrated of the same order (Gujarati 1995). For instance, time series are integrated of order 1, denoted I(1), if differencing the nonstationary time series once yields stationary or I(0) time series.

[5] Both test statistics follow a nonstandard distribution, and critical values are listed in Osterwald-Lenum (1992).

[6] Due to a typical slaughter yield of about 74%, the lean price is generally computed as the live price divided by 0.74 (Wellman 1996).

[7] Models were also estimated in simple form following equation (2) and no difference in results were detected.