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Impact of Mandatory Price Reporting on Hog Market Integration

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This research examines whether mandatory price reporting (MPR) impacted price relationships among U.S. hog markets. Markets are cointegrated before and after MPR enactment, but not fully integrated in either period. Terminal markets adjust to shocks in the Iowa-Southern Minnesota market more quickly and Iowa-Southern Minnesota prices adjust to shocks in terminal markets more slowly following MPR enactment. Granger causality tests indicate a causal flow from terminal markets to Iowa-Southern Minnesota prices before MPR and a causal reversal after MPR enactment. These results likely reflect decreases in volume of negotiated sales, particularly in terminal markets, and greater reliance on mandatorily reported prices for market information.

Key Words: cointegration, hog markets, mandatory price reporting, market integration, regime shift

JEL Classification: Q13

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 publicly available by the U.S. Department of Agricultural's Agricultural Marketing Service (2009) became increasingly scrutinized as being unreliable or unrepresentative of industry trade. Congress passed the Livestock Mandatory Reporting Act of 1999 (Federal Register, 2000), which went into effect in April 2001, with an aim to facilitate transparent price discovery and encourage competition (Azzam, 2003; Pendell and Schroeder, 2006). The law requires federally-inspected pork plants that slaughter over 100,000 head annually to report daily prices, volumes, and terms of sale for domestic hog purchases from the previous business day.

The objective of this study is to examine whether mandatory price reporting (MPR) has influenced spatial market integration among four domestic spot markets for hogs and also the degree of integration with futures market prices. Three terminal markets (Peoria, Illinois; St. Joseph, Missouri; and St. Paul, Minnesota) and one regional market (Iowa-Southern Minnesota) are considered.[1] With more complete regional price and transaction data available following

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[1] A terminal market refers to a market (e.g., buyer or set of buyers) at a specific location where various quantities of a commodity are pooled from multiple sellers on a large scale for either transfer or processing. Regional market prices reflect prices reported at various locations within the region.

MPR, integration and speed of price adjustment between spatially dispersed markets may be expected to increase. However, the implications for specific terminal markets are less clear.

There exists a rich body of literature on spatial market integration in agriculture with empirical applications extending from crops (e.g., Brester and Goodwin, 1993; Franken et al., 2005; Goodwin, 1992a; Goodwin and Piggott, 2001) to cattle (e.g., Pendell and Schroeder, 2006; Schroeder and Goodwin, 1990) to hogs (e.g., Benson et al., 1994; Schroeder and Goodwin, 1991). In fed cattle markets, there is mixed evidence on whether mandatory price reporting offers more transparency than voluntary systems (Fausti and Diersen, 2004; Fausti et al., 2010). Pendell and Schroeder (2006) find that fed cattle markets became more fully integrated following enactment of MPR. To date, no study has examined price integration among U.S. hog markets, let alone whether MPR has had an effect.

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., Carter and Mohapatra, 2008; Yang, Bessler, and Leatham, 2001). In general, these studies suggest that futures lead spot prices, which is consistent with market efficiency theory. Cointegration of spot and futures hog prices is revisited in the current study, and impacts of MPR on spatial price relationships are investigated.

Weekly average prices for four U.S. spot markets and nearby hog futures contracts for 1992 through August 2009 are analyzed.[2] Conventional methods for investigating market integration are employed due to a lack of trade (i.e., transfer and transfer cost) data enabling more sophisticated techniques, while recognizing that empirical findings of these methods may not always be indicative of the degree of market efficiency (i.e., Barrett and Li, 2002; McNew and Fackler, 1997).[3] 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. Vector Error Correction speed-of-adjustment coefficients before and after MPR are examined to ascertain whether the policy change has affected how quickly markets return to long-run equilibrium following price shocks. 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., Benson et al., 1994; Franken et al., 2005; Goodwin, 1992a,b; Goodwin and Piggott, 2001; Pendell and Schroeder, 2006; Schroeder and Goodwin, 1990).[4] This section reviews selected studies and emphasizes research on livestock industries with mandatory price reporting, specifically, cattle and hog markets.

Several studies of spatial price relationships, including the most comprehensive study on the effects of MPR to date (i.e., Pendell and Schroeder, 2006), have been conducted on cattle markets. Fed cattle price relationships during the 1970s

[2] As the mandatorily reported Iowa-Southern Minnesota price series of primary interest is available back to 1992 only as weekly averages, the analysis is conducted on weekly average prices which potentially introduce autocorrelation that mid-week closing prices would not (Working, 1960).

[3] McNew and Fackler (1997) note that if arbitrage (i.e., transport and other transfer) costs are nonstationary or if prices wander in periods with no trade, then cointegration may not be found even in well-functioning markets. Noting similar complications with cointegration, error correction, and Granger causality methods, Barrett and Li (2002) offer a mixture distribution model utilizing price, transfer cost, and trade flow data to distinguish between perfect integration and segmented equilibrium (both consistent with spatial equilibrium) and imperfect integration and segmented disequilibrium (both inconsistent with spatial equilibrium).

[4] Markets perform efficiently when they are integrated (i.e., when the difference in prices at two locations is less than or equal to transportation and other transfer costs). "The equilibrium condition binds with equality when trade occurs" (Barrett and Li, 2002, p. 293).

and 1980s were examined by Bailey and Brorsen (1985), Koontz, Garcia, and Hudson (1990), and Schroeder and Goodwin (1990), who found larger volume and direct 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 quicker than smaller markets (Schroeder and Goodwin, 1990). Goodwin and Schroeder (1991) found limited cointegration, especially among geographically dispersed fed cattle markets, which increased over time with concentration in cattle slaughtering in the 1980s, possibly reflecting decreasing trade and information costs or noncompetitive basing-point pricing. Goodwin (1992b) confirmed a gradual structural change toward subsequently faster price adjustment from the mid-1970s through early 1980s, and Schroeder (1997) found faster speed of adjustment in early 1990s fed cattle for processing plants in close proximity and slower speed of adjustment for larger plants and plants with fewer cash transactions.

Fausti and Diersen (2004) examined the relationship between fed cattle prices reported voluntarily for Nebraska direct and mandatorily for South Dakota, and concluded that voluntary price reporting was as transparent for price discovery as the mandatory system. In contrast, Fausti et al. (2010) found evidence that MPR enhanced transparency of publicly reported fed cattle grid premiums and discounts, and Pendell and Schroeder's (2006) regime shift cointegration model indicated that cattle markets became more fully integrated following enactment of MPR. Ward (2008) found for both cattle and hogs that cash prices lead alternative marketing arrangements (AMAs) in rising markets and trail them in declining markets. Lee, Ward, and Brorsen (2010) found that cash and AMAs prices are cointegrated with all but one procurement method. While bidirectional causality was found among some procurement prices, cash prices Granger cause all other procurement prices, indicating that cash markets remain of central importance to price discovery.[5]

In general, few hog price integration studies have examined spatial relationships (e.g., Benson et al., 1994; Chen and Lee, 2008; Faminow and Benson, 1990), as the focus has been on relationships between cash and futures prices (e.g., Carter and Mohapatra, 2008; Schroeder and Goodwin, 1991; Yang, Bessler, and Leatham, 2001), and as discussed above, cash and AMA prices (e.g., Lee, Ward, and Brorsen, 2010; Ward, 2008). Cash and futures hog prices generally were cointegrated (Carter and Mohapatra, 2008; Yang, Bessler, and Leatham, 2001),[6] and price discovery originated in the futures market (Carter and Mohapatra, 2008; Schroeder and Goodwin, 1991) with futures being fairly unbiased predictors of future cash prices except for distant contracts (Carter and Mohapatra, 2008; Yang, Bessler, and Leatham, 2001). Studies of spatial hog price relationships found evidence of inefficiencies (i.e., possible basing-point pricing) among Canadian markets (Faminow and Benson, 1990), cointegration among Canadian prices and the U.S. price (Benson et al., 1994), and integration among markets in Taiwan (Chen and Lee, 2008). Importantly, none of these studies examined hog price integration among multiple U.S. markets.

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 (2006). 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

[5] 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.

[6] An earlier study by Schroeder and Goodwin (1991) found a lack of cointegration between Omaha cash prices and Chicago Mercantile Exchange futures prices for hogs.

equilibrium relationships may be estimated.[7] The well-known test for cointegration between two spatial markets attributed to Engle and Granger (1987) is estimated by ordinary least squares (OLS) as:

$$\text{(1)} \quad \textit{Model I, Standard Cointegration:} \quad Y_t = \alpha_0 + \alpha_1 Z_t + e_t,$$

where Y_t and Z_t are individual nonstationary 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).[8] 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:

$$\text{(2)} \quad \textit{Model II, Regime Shift Cointegration:} \quad Y_t = \alpha_0 + \alpha_1 D_t + \alpha_2 Z_t + \alpha_3 Z_t D_t + e_t,$$

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 enactment of MPR; α_2 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. Notably, other changes occurring in the hog industry over the time period studied (e.g., declining volumes of hogs in spot markets) are not accounted for in this model due mostly to a lack of adequate data.

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:

$$\text{(3)} \quad \begin{aligned} \Delta P_t = {} & \alpha_0 + \alpha_1 D_t + \alpha_2 Z_t + \alpha_3 Z_t D_t \\ & + \sum_{k=1}^{k} \beta_{11}(k)\, \Delta P_{t-k} \\ & + \sum_{k=1}^{k} \beta_{12}(k)\, \Delta Z_{t-k} + \Omega t, \end{aligned}$$

[7] 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. While research suggests that nominal commodity spot prices do not often possess unit roots (i.e., prices are stationary) and findings of nonstationarity are sensitive to specification of the data generating process (c.f., Wang and Tomek, 2007), these procedures work relatively well in empirical work.

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

where t refers to time ($t = 1, 2, \ldots, T$), which for this study is weeks; k is the lag length; and Ω is a $n \times 1$ vector of normally distributed random errors. The specification of Equation (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 price relationships between spot market locations, error correction VAR, or Vector Error Correction (VEC) models, incorporating the binary MPR dummy D_t 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:

$$\begin{aligned} \Delta P_t = {} & \beta_0 + \beta_1 \hat{e}_{t-1} + \beta_2 (\hat{e}_{t-1} \times D_t) \\ & + \sum_{k=1}^{K} \beta_{11}(k)\, \Delta P_{t-k} \\ & + \sum_{k=1}^{K} \beta_{12}(k)\, \Delta Z_{t-k} + \lambda_t, \end{aligned} \tag{4}$$

where variables and subscripts are as defined in Equation (3), and λ is a $n \times 1$ vector of normally distributed random errors. In Equation (4), β_1 measures the speed-of-adjustment or the one period lagged errors' effect on a relative price change for the entire sample period, and β_2 measures the change in the magnitude of the speed-of-adjustment for a relative price change during only the time period that MPR is in effect. This effect is captured using an interaction term specified as the product of one period lagged errors and the binary dummy variable D_t equaling one in the MPR period. The lagged error terms specified in Equation (4) are obtained from the OLS estimation of Equation (1). The next two terms are lagged price change variables following from the standard VEC model. A speed-of-adjustment coefficient (β_1) 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 (an adjusted or aggregate) speed-of-adjustment ($\beta_1 + \beta_2$) nearer to one in absolute value should be observed.

Data

Weekly average prices are analyzed from 1992 through August 2009 for four U.S. spot markets and nearby hog futures contracts rolled over 1 week prior to expiration.[9] Voluntarily reported U.S. spot prices for terminal markets in St. Joseph, Missouri; Peoria, Illinois; and St. Paul, Minnesota are obtained from Plain (2010).[10] Mandatorily reported Iowa-Southern Minnesota interior market prices are obtained from the Livestock Market Information Center. Since MPR applies to hog buyers purchasing over 100,000 head of barrows/gilts annually, the Iowa-Southern Minnesota series reflects purchases made by qualifying hog buyers from across the north central region, but reported prices for the St. Paul terminal market are determined by all transactions occurring at that location. Nearby Chicago Mercantile Exchange (CME) lean hog futures prices are obtained from the Commodity Research Bureau.

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). The former contract relied on physical delivery, while the latter is cash settled based on the CME Lean Hog Index (Frank et al., 2008). The rationale for these changes was to expand the contract's usefulness as a hedging instrument to domestic and international hog producers and packers as well as to hog and pork importers and exporters. With implementation of MPR in April 2001, Iowa-Southern Minnesota and other markets began reporting prices on a carcass basis. Livestock

[9] In the empirical analysis, dummy variables are used to account for futures contract roll-over. Specifically, a different dummy variable is included for every rollover and for every time a rollover occurs in a lagged variable.

[10] Upon closure of St. Paul location in April 2008, participants in that market initiated business in Zumbrota, Minnesota. Hence, the St. Paul series is constructed of St. Paul prices prior to April 15, 2008 and Zumbrota prices thereafter.

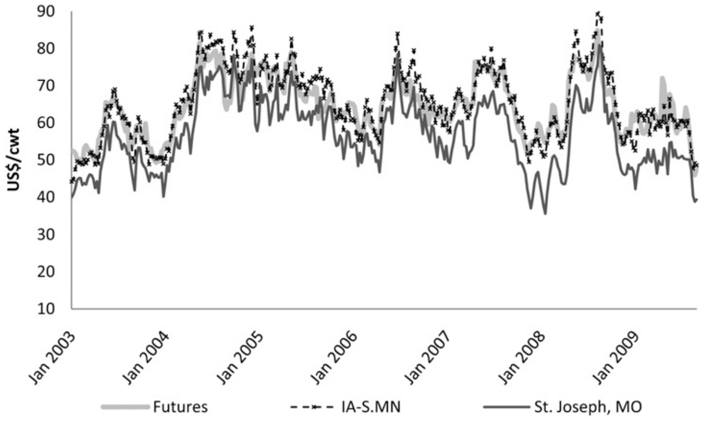


Figure 1. U.S. Hog Carcass Prices, 1992–2009

Market Information Center 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.[11] Lean value adjusted prices for only the futures, Iowa-Southern Minnesota, and St. Joseph markets are graphed in Figure 1 as representative markets that, like those not shown, exhibit fairly similar patterns. Summary statistics and correlations are reported for hog prices on a carcass basis (Table 1). The correlations among prices are all above 0.90 with the exception of Peoria's correlation with CME and Iowa-Southern Minnesota (IAMN), which is 0.89 in both cases.

Results

Prior to market integration analysis, Augmented Dickey-Fuller tests of nonstationarity were performed. The appropriate lag structure for the ADF tests and all subsequent models was determined by minimizing the Akaike Information Criteria, and thereby was set to four lags. In all cases, the null hypothesis of nonstationarity could not be rejected at the 5% significance 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 5% 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 otherwise) 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 ($\alpha_2 + \alpha_3 = 1$).[12] This finding is largely consistent across alternative

[11] 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).

[12] Models were also estimated following Equation (2) and no difference in results was detected.

Table 1. Summary Statistics and Correlations for Futures and Spot Market Hog Prices ($/cwt)

	Summary Statistics				Correlations				
	Mean	Max	Min	SD	CME	IAMN	St. Paul	St. Joseph	Peoria
CME[a]	61.47	87.68	23.03	10.44	1.00				
IAMN[b]	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

Source: Nearby Chicago Mercantile Exchange lean hog futures prices are obtained from the Commodity Research Bureau, Iowa-Southern Minnesota interior hog market prices are obtained from the Livestock Market Information Center, and hog prices for terminal markets in St. Joseph, Missouri; Peoria, Illinois; and St. Paul, Minnesota are obtained from University of Missouri Extension specialist, Ron Plain.

Note: N = 922 observations.

[a] CME denotes the nearby Chicago Mercantile Exchange group lean hog futures price series.

[b] IAMN denotes the Iowa-Southern Minnesota hog price series.

orderings of dependent and independent variables.[13] Hence, we present 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 the VEC model 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. Specifically, we compare how quickly the terminal price series (i.e., St. Joseph, Missouri; Peoria, Illinois; and St. Paul, Minnesota) adjusts to price shocks in the IAMN market with how quickly the IAMN price series adjusts to price shocks in one of the terminal markets and examine how MPR has impacted these relationships.

Recall, if MPR improves availability of reliable price information, then the adjusted or aggregate speed-of-adjustment measure ($\beta_1 + \beta_2$) should be nearer to one in absolute value than the simple, unadjusted measure (β_1). As an example, consider the results for St. Joseph/IAMN and IAMN/St. Joseph market pairs in Table 4. Notice that speed-of-adjustment coefficients (β_1) indicate that, in response to a unit deviation from equilibrium in period t – 1, the IAMN price falls 0.0532 units and the St. Joseph price rises 0.0536 units within 1 week, quickly converging toward long-run equilibrium. While the absolute values of β_1 generally are similar in magnitude, regardless of the ordering of pairs, differential impacts are implied for MPR by β_2 coefficients, which measure the change in magnitude of the speed-of-adjustment after MPR enactment. For instance, β_2 is over twice as large in absolute value for St. Joseph/IAMN than for IAMN/St. Joseph (-0.1698 and 0.0643, respectively). Consequently, comparing the net impact ($\beta_1 + \beta_2$) with the speed-of-adjustment coefficient (β_1) reveals faster adjustment of St. Joseph prices to price shocks in IAMN ($|-0.1162| > |0.0536|$) and slower adjustment of IAMN prices to price shocks in St. Joseph ($|0.0112| < |-0.0532|$) in the MPR period. This observation holds for other pairings of terminal markets with IAMN as well. Hence, terminal markets adjust to shocks in the IAMN market more quickly and IAMN prices adjust to shocks in terminal markets more slowly following MPR enactment. This finding likely reflects increased confidence in the mandatorily reported IAMN price as a representative source of market information, but also may reflect the relative size of these markets. Given the decreasing volume of terminal markets, price movements in the larger IAMN market evolve somewhat independently of events in terminal markets, but

[13] Price discovery could occur simultaneously in multiple markets, and hence, price integration relationships among market pairs should be considered in both possible *causal* directions.

Table 2. Summary of Johansen Unrestricted Cointegration Rank Test Statistics

	Before MPR		After MPR	
Market Pairs	Eigenvalue	Trace Statistic	Eigenvalue	Trace Statistic
IAMN[a]/Peoria	0.016**	17.118	0.024**	18.637
IAMN/St. Joseph	0.016**	19.104	0.040**	26.038
IAMN/St. Paul	0.016**	18.934	0.035**	23.698
IAMN/CME[b]	0.054**	65.731	0.091**	50.275
St. Joseph/Peoria	0.018**	30.489	0.032**	24.355
St. Joseph/St. Paul	0.071**	82.618	0.071**	41.644
St. Joseph/CME	0.048**	60.119	0.063**	37.249
St. Paul/Peoria	0.018**	29.503	0.037**	25.980
St. Paul/CME	0.045**	58.596	0.054**	33.717
Peoria/CME	0.026**	38.911	0.037**	26.040

Notes: 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.
[a] IAMN denotes the Iowa-Southern Minnesota hog price series.
[b] CME denotes the Chicago Mercantile Exchange Group lean hog futures price series.

terminal markets respond quickly to changes in the IAMN price.[14]

Table 5 presents the results of Granger causality tests corresponding to the VEC framework to determine the extent to which lagged prices for one market influence prices in another market. Test statistics for the null hypothesis of no causality are presented for portions of the sample before and after enactment of MPR, as well as the entire sample period. For the entire period, the null hypothesis is rejected at conventional levels for nearly every case, with the exception that Peoria does not Granger cause IAMN. Comparison of results for pre-MPR and post-MPR periods offers more interesting insights. Prior to MPR, IAMN prices did not Granger cause terminal prices, but St. Paul and Peoria prices Granger caused IAMN prices. However, the causality relationships reverse in the period following enactment of MPR. Again, the results seem to reflect increased reliance on the mandatorily reported IAMN price series as a representative source of market information and less attention to price information in terminal markets where volumes are dwindling.

The IAMN/St. Joseph results are a curious case. St. Joseph Granger causes IAMN prices after but not prior to enactment of MPR, whereas the opposite result is observed for other terminal markets. Personal communication with Ron Plain (2010), University of Missouri Extension specialist, provides insight. Early in the sample period, hogs sold at St. Joseph were slaughtered in Missouri. With the closing of slaughtering plants in St. Joseph in December 1993 and in Marshall in July 2001 many of the hogs sold in St. Joseph were transported to slaughtering facilities in Iowa, particularly before Triumph opened a pork processing plant in St. Joseph in January 2006. These events likely contribute to the relatively higher causal impact of St. Joseph prices on IAMN prices in the period following MPR.

As mentioned above, changes in price relationships observed here may partly reflect declining volume in negotiated or spot markets. 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 Goodwin and Schroeder

[14] Based on an anonymous reviewer's suggestion that some of the slow adjustment observed may reflect non-zero intra-week adjustments not captured by weekly average prices, contemporaneous relations are investigated by applying Spirtes, Glymour, and Scheines' (2000) PC algorithm to innovations (i.e., residuals from multiple time series VEC models) following Stockton, Bessler, and Wilson (2010). Causal direction of innovations at the 0.05 significance level, as represented in directed acyclic graphs, suggest contemporaneous flows of information that are largely consistent with Granger causality findings presented below. Directed acyclic graphs are available from the authors upon request.

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

Dependent Market/ Independent Market[a]	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)
IAMN[b]/CME[c]	0.1595	0.4577***	−1.4728	0.0343*	0.0000	0.0000
	(0.8385)	(0.0269)	(1.3192)	(0.0194)		
St. Joseph/CME	−0.0522	0.4540***	−1.9980	0.0158	0.0000	0.0000
	(0.8803)	(0.0283)	(1.3850)	(0.0204)		
St. Paul/CME	0.8120	0.5037***	−2.2003	0.0337	0.0000	0.0000
	(0.9179)	(0.0295)	(1.4441)	(0.0213)		
Peoria/CME	0.0141	0.4456***	−0.9621	0.0101	0.0000	0.0000
	(0.8209)	(0.0264)	(1.2916)	(0.0190)		
St. Joseph/IAMN	0.0974	0.9107***	−1.0063	−0.0098	2.3100×10^{-10}	2.4070×10^{-08}
	(0.4771)	(0.0173)	(0.7105)	(0.0104)		
St. Paul/IAMN	0.8556*	0.9922***	−0.6235	−0.0025	0.6384	0.5183
	(0.4566)	(0.0165)	(0.6801)	(0.0100)		
Peoria/IAMN	−0.0086	0.8683***	0.4416	−0.0223**	3.3306×10^{-16}	0.0000
	(0.4469)	(0.0162)	(0.6657)	(0.0098)		
St. Joseph/St. Paul	−1.0309**	0.8904***	0.4166	−0.0235**	7.1054×10^{-15}	0.0000
	(0.4344)	(0.0141)	(0.6314)	(0.0096)		
Peoria/St. Paul	−1.0106***	0.8525***	1.603***	−0.0324***	0.0000	0.0000
	(0.3838)	(0.0125)	(0.5579)	(0.0085)		
St. Joseph/Peoria	−0.0341	0.9504***	−0.7582	0.0006	0.0023	0.0025
	(0.4468)	(0.0163)	(0.6491)	(0.0100)		

Note: N = 922.

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

[a] Ordering of dependent and independent variables follows from Equation (2). Overall, findings do not qualitatively change when dependent and independent variable are switched (i.e., the dependent variable moves from the left-hand-side to the right-hand-side and the independent variable moves from the right-hand-side to the left-hand-side).

[b] IAMN denotes the Iowa-Southern Minnesota hog price series.

[c] CME denotes the Chicago Mercantile Exchange Group lean hog futures price series.

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

	VEC 3		
Dependent Market/ Independent Market[a]	Speed-of-Adjustment Coefficient (entire period) (β_1)	Size of Speed-of-Adjustment after MPR (β_2)	Net Impact ($=\beta_1 + \beta_2$)
St. Joseph/IAMN[b]	0.0536	−0.1698**	−0.1162
	(0.0499)	(0.0847)	
St. Paul/IAMN	0.0218	−0.1076	−0.0858
	(0.0552)	(0.0911)	
Peoria/IAMN	0.0202	−0.0750	−0.0549
	(0.0381)	(0.0618)	
IAMN/St. Joseph	−0.0532	0.0643	0.0112
	(0.0563)	(0.0897)	
IAMN/St. Paul	−0.0257	0.0216	−0.0041
	(0.0574)	(0.0911)	
IAMN/Peoria	−0.0381	0.0338	−0.0043
	(0.0441)	(0.0665)	

Note: N = 922.
***, **, * denote statistical significance at 1%, 5%, 10% levels, respectively. Standard errors are in parentheses. Lag length is set to 4.
[a] Ordering of dependent and independent variables follows from Equation (4).
[b] IAMN denotes the Iowa-Southern Minnesota hog price series.

(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, and the adjusted R^2 is quite small.

Conclusions

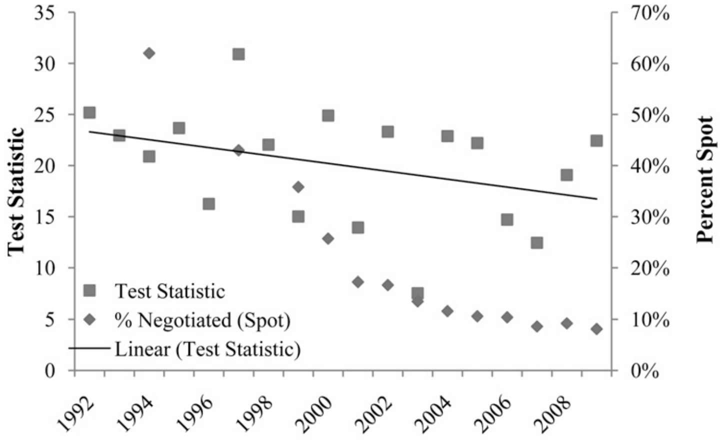
This paper investigates price relationships among various U.S. hog markets and whether mandatory price reporting, which is intended to facilitate transparent price discovery, has had any detectable impact on these price relationships. Previous research on the impacts of MPR in cattle markets is mixed. 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

Table 5. Granger Causality for Hog Prices in Vector Error Correction Models

Dependent Market/ Independent Market[a]	χ^2 Test Statistic: Pre-MPR	Post-MPR	Entire Period
St. Joseph/IAMN[b]	7.59	80.41***	131.93***
St. Paul/IAMN	9.03*	15.52***	18.35***
Peoria/IAMN	5.39	37.74***	46.67***
IAMN/St. Joseph	6.44	10.09**	11.01**
IAMN/St. Paul	26.35***	4.87	15.72***
IAMN/Peoria	13.00**	5.28	2.42
N	483	439	922

***, **, * denote statistical significance at 1%, 5%, 10% levels, respectively. Lag length is set to 4.
[a] Ordering of dependent and independent variables follows from Equation (4).
[b] IAMN denotes the Iowa-Southern Minnesota hog price series.



Note(s): Bootstrapped regression: Test Statistic = 16.6390 + 12.4451×Percent Spot + error, Adjusted R^2 = 0.0267
(3.5924) (25.4523)

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

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-Southern Minnesota regional price in both periods, meaning that the former responds one-for-one to changes in the latter. Lack of full integration may reflect transaction costs and require threshold analysis beyond the current scope of this research (Goodwin and Piggott, 2001; Goodwin and Schroeder, 1991).

Vector error correction models indicate that terminal hog markets adjust to price shocks in larger Iowa-Southern Minnesota markets more quickly and Iowa-Southern Minnesota prices adjust to shocks in smaller terminal markets more slowly following enactment of MPR. Furthermore, Granger causality tests generally indicate that terminal prices Granger cause Iowa-Southern Minnesota prices prior to but not after the enactment of MPR, and Iowa-Southern Minnesota prices Granger cause terminal prices after but not prior to enactment of MPR. These results may reflect a number of other events coinciding with enactment of MPR. Mandatory price reporting of certain negotiated prices, as well as prices of alternative marketing arrangements, may have shifted focus to these measures as sources of reliable market information. When comparing speed-of-adjustment coefficients for negotiated and alternative marketing arrangement prices, Lee, Ward, and Brorsen (2010) interpret results as indicating that swine formula prices primarily adjust to divergences with negotiated prices but negotiated prices do more of the adjusting to divergences with other formula prices and other purchase prices. Their Granger causality tests also indicate that negotiated cash prices unidirectionally cause other formula prices and other purchases prices, but, there is relatively strong feedback from swine formula prices. Additionally, the proportion of hog sales transacted through spot market exchanges, particularly in terminal markets, 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.

As mentioned earlier, the conventional methods for investigating market price integration employed here are less precise indicators of market efficiency than more sophisticated techniques requiring comprehensive trade (i.e., price, transfer, and transfer cost) data (c.f., Barrett and Li, 2002; McNew and Fackler, 1997). Given the limitations of the data and methods employed here, the results should be cautiously viewed as *prima facie* evidence on the degree of market integration (c.f., Bessler and Covey, 1993). Future research may yield more definitive results if more complete and higher frequency trade data can be obtained.

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References

Azzam, A. "Market Transparency and Market Structure: The Livestock Mandatory Reporting Act of 1999." *American Journal of Agricultural Economics* 85,2(2003):387–95.

Bailey, D., and B.W. Brorsen. "Dynamics of Regional Fed Cattle Prices." *Western Journal of Agricultural Economics* 10,1(1985):126–33.

Barrett, C.B., and J.R. Li. "Distinguishing Between Equilibrium and Integration in Spatial Price Analysis." *American Journal of Agricultural Economics* 84,2(2002):292–307.

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

Bessler, D.A., and T. Covey. "On the Search for Econometric Structure in Agriculture." *American Journal of Agricultural Economics* 75(1993): 41–47.

Brester, G.W., and B.K. Goodwin. "Vertical and Horizontal Price Linkages and Market Concentration of the U.S. Wheat Milling Industry." *Review of Agricultural Economics* 15,3(1993): 507–19.

Carter, C.A., and S. Mohapatra. "How Reliable are Hog Futures as Forecasts?" *American Journal of Agricultural Economics* 90,2(2008): 367–78.

Chen, P.-F., and C.-C. Lee. "Nonlinear Adjustments in Deviations from the Law of One Price for Wholesale Hog Prices." *Agricultural Economics* 39,1(2008):123–34.

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

Engle, R.F., and C.W.J. Granger. "Co-integration and Error Correction: Representation, Estimation, and Testing." *Econometrica* 55(1987):251–76.

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

Federal Register. Part V, Department of Agriculture, Agricultural Marketing Service, Livestock and Grain Market News Branch: Livestock Mandatory Reporting; Final Rule. National Archives and Records Administration. Friday, December 1, 2000.

Fausti, S.W., and M.A. Diersen. "The Voluntary Reporting System's Ability to Provide Price Transparency in the Cash Market for Dressed Steers: Evidence from South Dakota." *Journal of Agricultural and Resource Economics* 29,3 (2004):553–66.

Fausti, S.W., B.A. Qasmi, J. Li, and M.A. Diersen. "The Effect of the Livestock Mandatory Reporting Act on Market Transparency and Grid Price Dispersion." *Agricultural and Resource Economics Review* 39,3(2010):457–67.

Frank, J., M. Gómez, E. Kunda, and P. Garcia. "Cash Settlement of Lean Hog Futures Contracts Reexamined." Paper presented at the NCCC-134 Conference on Applied Commodity Price Analysis, Forecasting, and Market Risk Management, St. Louis, Missouri, April 21–22, 2008.

Franken, J.R.V., J.L. Parcell, M.E. Sykuta, and C.F. Fulcher. "Market Integration: Case Studies of Structural Change." *Agricultural and Resource Economics Review* 43,2(2005):163–72.

Goodwin, B. "Multivariate Cointegration Tests and the Law of One Price in International Wheat Markets." *Review of Agricultural Economics* 14,1(1992a):117–24.

———. "Forecasting Cattle Prices in the Presence of Structural Change." *Southern Journal of Agricultural Economics* 24,2(1992b):11–22.

Goodwin, B., and N. Piggott. "Spatial Market Integration in the Presence of Threshold Effects." *American Journal of Agricultural Economics* 83,2(2001):302–17.

Goodwin, B.K., and T.C. Schroeder. "Cointegration Tests and Spatial Price Linkages in Regional Cattle Markets." *American Journal of Agricultural Economics* 73(1991):453–64.

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

Grimes, G., and R. Plain. *U.S. Hog Marketing Contract Study*. Columbia, MO: University of Missouri Department of Agricultural Economics, 2005.

———. *U.S. Hog Marketing Contract Study*. Columbia, MO: University of Missouri Department of Agricultural Economics, 2007.

Gujarati, D. *Basic Econometrics*, 3rd ed. New York: McGraw-Hill, Inc, 1995.

Johansen, S. "Statistical Analysis of Cointegration Vectors." *Journal of Economic Dynamics & Control* 12(1988):213–54.

Koontz, S.R., P. Garcia, and M.A. Hudson. "Dominant-Satellite Relationships between Live Cattle Cash and Futures Markets." *Journal of Futures Markets* 10,2(1990):123–36.

Lee, Y., C.E. Ward, and B.W. Brorsen. "Relationships among Prices across Alternative Marketing Arrangements for Fed Cattle and Hogs." Paper presented at the Southern Agricultural Economics Association Annual Meeting, Orlando, FL, February 6–9, 2010.

McNew, K., and P.L. Fackler. "Testing Market Equilibrium: Is Cointegration Informative." *Journal of Agricultural and Resource Economics* 22,2(1997):191–207.

Osterwald-Lenum, M. "A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics." *Oxford Bulletin of Economics and Statistics* 54(1992):461–71.

Pendell, D.L., and T.C. Schroeder. "Impact of Mandatory Price Reporting on Fed Cattle Market Integration." *Journal of Agricultural and Resource Economics* 31,3(2006):568–79.

Plain, R. Personal Communication. Department of Agricultural and Applied Economics, University of Missouri, February 15, 2010.

Schroeder, T.C. "Fed Cattle Spatial Transactions Price Relationships." *Journal of Agricultural and Applied Economics* 29,2(1997):347–62.

Schroeder, T.C., and B.K. Goodwin. "Regional Fed Cattle Price Dynamics." *Western Journal of Agricultural Economics* 15,1(1990):111–22.

———. "Price Discovery and Cointegration for Live Hogs." *Journal of Futures Markets* 11,6 (1991):685–96.

Spirtes, P., C. Glymour, and R. Scheines. *Causality, Prediction, and Search*. Cambridge, MA: MIT Press, 2000.

Stockton, M.C., D.A. Bessler, and R.K. Wilson. "Price Discovery in Nebraska Cattle Markets." *Journal of Agricultural and Applied Economics* 42,1(2010):1–14.

Wang, D., and W.G. Tomek. "Commodity Prices and Unit Root Tests." *American Journal of Agricultural Economics* 89,4(2007):873–89.

Ward, C.E. "Preferential Cattle and Hog Pricing by Packers: Evidence from Mandatory Price Reports." Paper presented at the Western Agricultural Economics Association Annual Meeting, Big Sky, MT, June 23–24, 2008. Internet site: http://ageconsearch.umn.edu/bitstream/37989/2/AMAPrefP%27gWAEA.pdf (Accessed February 22, 2010).

Wellman, A. *Highlights of the New Chicago Mercantile Exchange Lean Hogs Futures Contract*. University of Nebraska Extension Publication NF96-261, 1996. Internet site: http://digitalcommons.unl.edu/cgi/viewcontent.cgi?article=1392&context=extensionhist (Accessed March 22, 2010).

Working, H. "Note on the Correlation of First Differences of Averages in a Random Chain." *Econometrica* 28,4(1960):916–18.

Yang, J., D.A. Bessler, and D.J. Leatham. "Asset Storability and Price Discovery in Commodity Futures Markets: A New Look." *Journal of Futures Markets* 21,3(2001):279–300.