Does the Boxed Beef Price Inform the Live Cattle Futures Price?

Kishore Joseph, Philip Garcia, Paul E. Peterson*

*Kishore Joseph (kjoseph5@csuchico.edu) is an Assistant Professor in Agricultural Business in the College of Agriculture at California State University Chico. Philip Garcia is the Thomas A. Hieronymus Distinguished Chair in Futures Markets and Director, Office of Futures and Options Research in the University of Illinois at Urbana-Champaign. Paul E. Peterson is the Clearing Corporation Foundation Clinical Professor of Derivatives Trading, in the Department of Agricultural and Consumer Economics at University of Illinois Urbana-Champaign.

Selected Paper Prepared for presentation at the 2016 Agricultural & Applied Economics Association Annual Meeting, Boston, Massachusetts, July 31-August 2

Abstract

We study the lead-lag relationship between live cattle futures and negotiated boxed beef cutout price. To account for temporal differences in the information content, Friday afternoon boxed beef cutout price are compared to both current day and one-day prior live cattle futures settlement prices. Extensive testing and innovation accounting based on VECM residuals indicate that the futures price leads the cutout price as the dominant source of information in the fed cattle market. The futures price has a strong predictive influence on the boxed beef cutout price and appears to assimilate fed cattle price information quicker than both contemporaneous and one-day ahead boxed beef cutout price. Newly-developed price discovery metrics interpreted to allow for a maximum cutout effect in the pricing process still identify the dominance of the current futures price, and nearly equal weighting for the lagged one-day futures price.

Introduction

In recent decades, the U.S. cattle industry has experienced increased concentration. These structural changes have fueled the debate on economic and policy issues related to price discovery in the cattle markets (Koontz and Ward 2011). Effective price discovery is critical to the efficient pricing of a commodity. Much of this debate has focused on the decline in the volume of negotiated fed cattle cash market transactions in the cash market, which could reduce the representativeness of these prices, lead to market manipulation and other distortions, and lower the quality of pricing. In this context, we evaluate the Chicago Mercantile Exchange (CME) live cattle futures price and the United States Department of Agriculture-Agricultural Marketing Service (USDA-AMS) boxed beef cutout price as competing sources of price information in the fed cattle market.

The Livestock Mandatory Reporting Act (LMRA) was passed by the Congress in 1999 to improve price transparency in livestock markets. Research suggests that there have been improvements in the quantity and quality of information available since the implementation of the LMRA in April 2001 (e.g., Perry et al. 2005; Ward 2006; Broyer and Brorsen 2013). Mandatory price reporting (MPR) has also raised concerns about the potential for coordination among beef packers (Wachenheim and Devuyst 2001; Azzam 2003; Njoroge 2003; Njoroje et al. 2007; Cai, Stiegert, and Koontz 2011a, 2011b). Regardless, the threat to price transparency exists as packers and producers increasingly choose non-negotiated cash methods to establish prices for the transfer of cattle ownership. In 2009 cash market transactions in which buyers and sellers negotiated the price and other terms of the transaction accounted for 50.4% of reported packer procurement, and formula pricing accounted for 36.5%. By 2013, negotiated transactions had declined to 29.4% of reported packer procurement and formula priced cattle volume had increased to 55.4%[1]. As these trends continue, the fed cattle market may be forced to look at alternate sources for timely and accurate price information.

While the fed cattle cash markets have been studied extensively, there has been less recent focus on the CME live cattle futures contract and on negotiated boxed beef cutout prices as tools for fed cattle price discovery. Live cattle futures quotes are readily available and are based on broad-based trading, making them valuable sources of information (Schroeder and Mintert 2000). Live cattle futures markets are also efficient in incorporating new supply and demand information (Garcia et al. 1988; McKenzie and Holt 2002). Prior research identifies the close link between cash and futures prices with the futures price leading cash price movements (Oellermann and Farris 1985; Koontz, Garcia, and Hudson 1990; Yang, Bessler, and Leatham 2002). Park, Jin, and Love (2011) argue that futures prices result from market information and do not drive decisions made along the beef supply chain. According to the authors, causality during the 1988-2005 periods is from the cash fed cattle price to the futures price. Using 2001-2010 data, Lee, Ward, and Brorsen (2012) report that negotiated cash prices led all alternative marketing agreement (AMA) prices except forward contracts which are closely tied to futures prices (Koontz, Garcia, and Hudson 1990). Ji and

[1]See USDA Grain Inspection, Packers and Stockyards Administration (GIPSA) 2013 report, table 17, at http://www.gipsa.usda.gov/psp/publication/ar/2013_ psp _ annual_ report.pdf

Chung (2012) report bidirectional causality between forward contract prices and cash fed cattle prices during the 2004-2011 periods.

The boxed beef cutout value, which reflects the composite price that packers receive from grocers, restaurants, processors, and others at the wholesale level for individual cuts of beef packaged in 40 to 60 pound boxes, is an alternate source of price discovery (Schroeder and Mintert 2000). The cutout price is closer to the retail sources and represents a broad-based average composed entirely of negotiated trades, making it a potentially viable alternative for pricing cattle. Reported boxed beef values represent a negotiated commitment from the packer to establish the price now and deliver the product within 21 calendar days. Boxed beef prices are reported by USDA so that market participants can use them in negotiating selling prices (Perry et al. 2005) and buying prices for beef. Packers also monitor these reports and use them as a benchmark for gauging their performance relative to competitors in the industry. As a result, boxed beef prices provide a reliable indication of price levels and changes at the next level in the cattle marketing chain.

Despite the informational value of the live cattle futures price and boxed beef cutouts in the fed cattle market, little recent attention has been given to the relationship between the two. Early studies by Ward (1981) identify the wholesale carcass price and nearby live cattle futures price as important variables in explaining the variation in the fed cattle price. Ward et al. (1997) suggest that the move towards a value-based cattle pricing system could shift the center of price discovery to the wholesale level. In contrast, Schroeder and Mintert (2000) show that the difference between boxed beef prices and cash fed cattle prices may vary depending on processing margins, making it difficult to use the cutout prices for fed cattle pricing.

The literature on the relationship between boxed beef and futures prices is limited by a lack of attention to the temporal aspects of boxed beef price reporting and by use of data not representative of the actual short-run pricing process in the market. The nearby futures price is the settlement price at 1:00 p.m., and the afternoon boxed beef value is the volume-weighted average of trades since 1:30 p.m. on the previous business day until 1:30 p.m. of the current business day. This implies that the futures settlement price may assimilate information from transactions underlying other contemporaneous cattle reports released by USDA.[2] A weekly frequency is appropriate in assessing the futures and cutout price relationship because the buyers and sellers of fed cattle operate in a window within a week where almost all transactions take place (Cai, Stiegert, and Koontz 2011b).

We empirically assess the relationship between the nearby CME live cattle futures price and the cutout value from the afternoon boxed beef report to find the leading indicator for fed cattle price movement. Consistent with the weekly fed cattle pricing process and to account for the potential of wholesale beef transaction information underlying a particular day's cutout price to be already available in same day futures price, the analysis is performed using the Friday cutout price from 1/19/2004 to 9/27/2013, and the corresponding Friday and previous day Thursday futures settlement prices. Friday cutout prices are selected because they are the most consistent price series available within the weekend, and represent the information that most producers have available to them going into the following week cash markets. Data are examined using time series procedures including cointegration and error correction modeling, testing for time-varying cointegration (Bierens and Martins 2010), innovation accounting, and newly-developed price discovery metrics (Yan and Zivot 2010; Putnins 2013).

Data

We use both Friday and Thursday futures settlement prices for the nearby CME live cattle futures contract from the Commodity Research Bureau (CRB) database. Futures contracts are rolled on the first day of the contract expiration month. The data are collected from

[2]MPR requires packers to submit the daily purchases of cattle that are scheduled for delivery 15-plus calendar days from the date they are purchased or priced as either a forward contract purchase or a formula marketing agreement purchase, respectively (Perry et al. 2005). These reports are generally released such that all trade information from 9:30 a.m. the previous day to 1:30 p.m. the current day becomes public by 3:00 p.m. the current day.

1/19/2004 to 9/27/2013. The live cattle electronic futures market opens at 9:05 a.m. on Monday and closes at 1:55 p.m. on Friday. The open outcry session opened at 9:05 a.m. and closed at 1 p.m. from Monday to Friday. The daily settlement price during the period is based solely on the trading activity in the pit between 12:59:30 p.m. and 1:00 p.m., and is reported shortly after 1:00 p.m.[3] The daily settlement price is the last futures price available before the afternoon boxed beef report is released at 3:00 p.m[4].

The cutout represents the estimated value of a beef carcass based on the sale prices received from packers for the individual beef items obtained from the carcass, and the USDA cutout formulation matches industry practices. During fabrication, carcasses are first broken into primal cuts and further fabricated into sub-primal cuts. A packer's overall cutout is based on the value and volume of sub-primal items being produced (USDA-AMS).[5] Boxed beef cutout values are released twice a day, at 11:00 a.m. and at 3:00 p.m.; we use the afternoon National Daily Boxed Beef Cutout and Boxed Beef Cuts-Negotiated Sales report (LM_ XB403) due to the thinness of the morning report. Cutout values are downloaded from USDA Livestock Market News Service historical data website (DATAMART).

The report provides data for two quality grades of beef, Choice and Select. The Choice cutout value is weighted 55% and the Select value is weighted 45% to match the par quality grade specification for the CME live cattle futures contract. Furthermore, because cutout values are on a carcass basis and the futures contract is on a live animal basis, we adjust the quality-weighted cutout value by the expected average hot yield of 63% (CME Rulebook, Chapter 101 Live Cattle Futures) and use this adjusted quality-weighted value as our cutout price. When the futures or the cutout price is not available on Fridays, both values are replaced with the next available weekday price going backward in time. We are careful

[3]See CME Rulebook, Chapter 8, Rule 813, for a detailed description of daily settlement procedures. Additional information is available at http://www.cmegroup.com/market-data/files/cme-group-settlement-procedures.pdf

[4]More recently, pit trading has ended and the electronic trading hours have changed for the CME live cattle futures contract.

[5]Detailed descriptions of the boxed beef report and calculation of the cutout values are available at http://www.ams.usda.gov/sites/default/files/media/USDADailyBoxedBeefReport.pdf.

to match prices from the same day when assessing the contemporaneous relationship, and maintain a one-day lag for the futures price when assessing the lagged relationship. All price series are expressed in cents per pound, and are converted into their natural logarithms consistent with procedures used in previous research.

Empirical Procedures

Multiple time series procedures are used to assess the relative information content of the live cattle futures price and cutout price. We rely primarily on the cointegration properties of the data to provide a rich characterization of the bivariate price relationship. We begin by testing for cointegration using the Johansen's (Johansen and Juselius 1990; Johansen 1992) procedure and then assess the stability in the price relationship using a time varying cointegration procedure proposed by Bierens and Martins (2010). The standard weak exogeneity tests are performed on cointegrated series followed by innovation accounting using forecast error variance decompositions and impulse response functions to identify the dynamic interaction of the two price series. Finally, newly developed price discovery metrics that rely on cointegration properties of the data are employed to identify the market providing the most relevant and timely information on fed cattle price.

Cointegration Analysis

In the presence of non-stationary series, the interaction between the futures price and the boxed beef price is examined by exploiting the cointegration relationship between them. A description of the Johansen's test for cointegration in the matrix form is given below. We start with the general P$^{\text{th}}$ order VAR model as follows

$$(1) \qquad \mathbf{\Delta Y_t} = \mathbf{D} + \mathbf{\Pi Y_{t-1}} + \sum_{j=1}^{p-1} \mathbf{\Gamma_j \Delta Y_{t-j}} + \boldsymbol{\varepsilon}_\mathbf{t} \quad (t = 1, ..., T),$$

where $\mathbf{Y_t}$ is an $(n \times 1)$ vector to be tested for cointegration, and $\mathbf{\Delta Y_t} = (\mathbf{Y_t} - \mathbf{Y_{t-1}})$; $\mathbf{D}$ is the deterministic term and may take different forms such as a vector of zeros or non-zero constants; $\mathbf{\Pi}$ and $\mathbf{\Gamma}$ are matrices of coefficients, $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$; and p is chosen so that $\boldsymbol{\varepsilon}_t$

is a multivariate normal white noise process with mean zero and finite covariance matrix ($\boldsymbol{\varepsilon_t} \backsim iid(0, \boldsymbol{\Sigma})$).

We begin by testing whether the vector $\mathbf{Y_t}$ is trend stationary rather than a multivariate unit root with drift process. Under the trend stationary hypothesis the matrix $\boldsymbol{\alpha\beta'}$ has full rank (k). One of Johansen's cointegration tests, the trace test, has this alternative hypothesis (Johansen and Juselius 1990; Johansen 1992). Since the Johansen's test is sensitive to the number of lags the Schwarz Bayesian Information Criterion (SBIC) is used to select the number of lags. The long-run pattern of price transmission is examined by testing the number of cointegration relations (r), using Trace tests and λ-max tests (Johansen and Juselius 1990; Johansen 1991). The cointegrating relationships explain the long-run equilibrium in prices, facilitated by the transmission of information. The rank of $\mathbf{\Pi}$ determines the number of cointegrating vectors, tested as follows

$$H(r) : \mathbf{\Pi} = \boldsymbol{\alpha\beta'}. \tag{2}$$

If prices are found to be cointegrated, a vector error correction model (VECM) imposing the cointegrating relationship is estimated to examine how prices adjust interactively under the constraint of the identified long-run equilibrium price relationships. The short-run dynamic pattern of price transmission can be observed from both, $\boldsymbol{\alpha}$ and $\mathbf{\Gamma_j}$, where $\boldsymbol{\alpha}$ parameter defines the short-run adjustments to the long-run relationship, and parameters $(\boldsymbol{\Gamma_j}, ...\boldsymbol{\Gamma_{p-1}})$ defines the short-run adjustment to changes in the process. Weak exogeneity tests for each price series $\mathbf{Y_t}$ (Johansen and Juselius 1990; Johansen 1992) are also performed as they allow us to identify the market that dominates in price discovery in the long run. This hypothesis is framed as

$$\boldsymbol{B'\alpha} = 0. \tag{3}$$

The null hypothesis is that each price does not respond to disturbances in the long-run relationship i.e., the i$^{\text{th}}$ row of the $\mathbf{\Pi}$ matrix is zero (Johansen and Juselius 1990, 1992; Johansen 1991).

In conventional cointegration analysis, cointegration vectors are assumed to be time-invariant. However the long-run relationship between prices may change due to structural breaks, and a time-invariant formulation of the cointegrating vector may no longer be appropriate (Hansen, 1992). A time-varying cointegration test based on Bierens and Martins (2010) is performed to assess if the relationship between futures and boxed beef prices has changed over the period. Bierens and Martins test the null hypothesis that the cointegrating vector $\boldsymbol{\beta}$ is constant against the alternative hypothesis that $\boldsymbol{\beta}$ is a function of time, $\boldsymbol{\beta_t}$. Specifically, the time-varying vector error correction model (TV-VECM) is

$$\mathbf{\Delta Y_t = D} + \boldsymbol{\alpha\beta'_t}\mathbf{Y_{t-1}} + \sum_{\mathbf{j=1}}^{\mathbf{p-1}} \mathbf{\Gamma_j \Delta Y_{t-j}} + \boldsymbol{\varepsilon_t}. \tag{4}$$

$\boldsymbol{\beta_t}$ is implemented by a series of Chebyshev time polynomials:

$$\boldsymbol{\beta_t} = \boldsymbol{\xi_0} + \boldsymbol{\xi_1}\mathbf{P_{1,T}(t)} + ... + \boldsymbol{\xi_m}\mathbf{P_{m,T}(t)}, \tag{5}$$

where $P_{i,T}(t)$ is a Chebyshev time polynomial of order i i.e.,

$$P_{0,T}(t) = 1, P_{i,T}(t) = \sqrt{2}cos(i\pi(t - .5)/T), \tag{6}$$

where $t = 1, 2, ..., T$, $i = 1, 2, 3, ...$, and ξ_i are the Fourier coefficients. The null hypothesis of time-invariant cointegration corresponds to the hypothesis that $H_0 : \xi_i = 0$ for $t = 1, 2, ..., m$. Under this null hypothesis the test statistic involved has a χ^2 distribution.

Innovation Accounting

We investigate the dynamic relationships between the series through innovation accounting, as recommended by Sims (1986) and Swanson and Granger (1997). Forecast error variance decompositions and impulse response functions are generated from the residuals of the VAR, or a VECM, to summarize the short-run dynamic linkages across various markets. Following Phillips (1998), we use an equivalent level VAR representation of the VECM imposing cointegration constraints and derive consistent results on forecast error variance decompositions and impulse response functions.

Price Discovery Measures

Finally, we employ two popular price discovery measures–Harris–McInish–Wood's Component Share (CS) and Hasbrouck's Information Share (IS), and use the specification by Yan and Zivot (2010) modified by Putnins (2013) to derive the Information Leadership Share (ILS) metric.[6] Both CS and IS rely on the notion that prices can deviate from each other in the short run due to market frictions, but their connection to the fundamental value will force them to converge in the long run (Lehman 2002). These measures are particularly useful when the price leadership between futures and cutout prices are assessed without involving cash price interactions. Similar to the classic Garbade and Silber (1983) price discovery approach, these price measures are based on an implicit unobservable efficient price common to all the underlying asset prices. The CS and IS price discovery measures, therefore, extend the Garbade and Silber (1983) formulation by allowing for cointegration between price series under the assumption of an underlying common random walk efficient price. The advantages of these metrics are that they can provide a measure of the proportional contribution to price discovery which VECM formulations are incapable of revealing.[7]

While the application of price discovery models beginning with the Garbade and Silber (1979) approach have been in arbitrage linked markets, the general principle outlined here does not hinge on the presence of standard arbitrage. Rather, we assume that information is assimilated and transferred between the futures and cutout markets when market participants trade in these markets with a rational view of the fundamental value of the asset. For instance, market participants are aware of average processing margins and would demand higher prices for cattle if wholesale prices keep rising. Similarly, packers would be unwilling to pay higher prices for cattle at lower wholesale prices. In short, an efficient price is discovered establishing a middle ground between buyer-seller outlook consistent with the prevailing

[6]The (CS) is often referred to as the Harris–McInish–Wood component share in literature because of their role in popularizing this measure. This measure is also known as the Permanent Transitory Model (PT) and is based on work by Gonzalo and Granger (1995).

[7]An application of the PT and IS measures to identify price discovery of floor and electronically traded corn, soybeans, and wheat futures contracts can be found in Martinez et al. (2011).

demand and supply situation, allowing farm and wholesale prices to move together in the long run. Information is incorporated into the futures market when market participants use the futures contract as a hedging tool.

The CS and IS differ in how they measure price discovery. Gonzalo and Granger (1995) divide a price vector into two parts, a permanent component interpreted as the implicit common efficient price (the driving force behind cointegration), and the temporary component reflecting the deviations due to market frictions. The authors show that the permanent component is a linear combination of all variables in the cointegrated system that can be easily estimated from a fully specified error correction model. The contribution of a price series to price discovery (CS) is then its normalized weight in the linear combination of prices that forms the common efficient price (Booth, So, and Tse 1999; Chu, Hsieh, and Tse 1999; Harris, McInish, and Wood 2002a).

In contrast, the common trend defined as the efficient price in Hasbrouck (1995) evolves as a random walk driven by the new information of an asset's value captured by the efficient price innovation represented in the Stock and Watson (1988) decomposition of a bivariate cointegrated series. Consider the Beveridge-Nelson (BN) decomposition (Bevereidge and Nelson 1981) of a cointegrated bivariate price vector with the level relationship adapted from Yan and Zivot (2010) below

$$\mathbf{P}_t = \mathbf{P}_0 + \mathbf{\Psi}(1)\sum_{j=1}^{t}\mathbf{e}_j + \mathbf{s}_t, \tag{7}$$

where $\mathbf{\Psi}(1) = \sum_{k=0}^{\infty}\mathbf{\Psi}_k$, $\mathbf{s}_t = (s_{1,t}, s_{2,t})' = \mathbf{\Psi}^*(L)e_t \sim I(0)$, and $\mathbf{\Psi}_k^* = -\sum_{j=k+1}^{\infty}\mathbf{\Psi}_j, k = 0, ..., \infty$. The matrix $\mathbf{\Psi}(1)$ contains the cumulative impacts of the innovations $\mathbf{e}_t$ on prices. As shown in Hasbrouck (1995), since $\boldsymbol{\beta}'\mathbf{\Psi}(1) = \mathbf{0}$ and $\boldsymbol{\beta} = (1, -1)'$, the rows of $\mathbf{\Psi}(1)$ are identical. Therefore, the long run impacts of an innovation $\mathbf{e}_t$ on each of the prices are identical. Representing $\mathbf{\Psi} = (\psi_1, \psi_2)'$ as the common row vector of $\mathbf{\Psi}(1)$, the permanent innovation can be defined as follows

$$\eta_t^p = \mathbf{\Psi}'\mathbf{e}_t = \psi_1 e_{1,t} + \psi_2 e_{2,t}. \tag{8}$$

Equation (8) can be represented using the common stochastic trend representation of Stock and Watson (1988) as follows

$$\mathbf{P}_t = \mathbf{P}_0 + \mathbf{1}m_t + \mathbf{s}_t, \tag{9}$$

where $\mathbf{1} = (1,1)', m_t = m_{t-1} + \eta_t^p$, and $\mathbf{s}_t = \mathbf{\Psi}^*(L)e_t$. The relationship indicates that each cointegrated price for the same underlying asset is composed of an unobservable common fundamental full-information value m_t, a transitory pricing error $s_{i,t}$ in market i, and a constant. This efficient price (m_t) evolves as a random walk driven by new information on the asset's future value captured by the efficient price innovation η_t^p. The pricing error $s_{i,t}$ captures any deviation of the price from its unobservable efficient price, and the remaining constant reflects any non-stochastic difference between the price and its efficient price. Since we are interested in how each series affects the efficient price innovation variance, Hasbrouck uses (8) to estimate the proportion of the variance of attributed to each series.

Hasbrouck (1995) proposes that the contribution to price discovery by a specific price, it's IS or information share, is the proportion of efficient price innovation variance that can be attributed to that price. When the price innovations across markets are correlated, the efficient price innovation variance is not attributed uniquely and the mean of the upper and lower bound of the estimated IS-values can be used.

The empirical representation of the two price discovery measures is straightforward. We follow Baillie et al. (2002) and calculate IS_1, IS_2, CS_1, and CS_2 from the error correction parameters and variance-covariance of the VECM error terms. Component shares are computed from the normalized orthogonal to the vector of error correction coefficients, $\boldsymbol{\alpha}_\perp = (\gamma_1, \gamma_2)'$, and are

$$CS_1 = \gamma_1 = \frac{\alpha_2}{(\alpha_2 - \alpha_1)}, CS_2 = \gamma_2 = \frac{\alpha_1}{(\alpha_1 - \alpha_2)}, \tag{10}$$

where α_1 and α_2 are the coefficients of the error correction term. A small (large) value CS is directly related to a small (large) contribution of market to the Gonzalo-Granger permanent component of prices (Booth, So, and Tse 1999; Chu, Hsieh, and Tse 1999; Harris, McInish, and Wood 200a). If $\alpha_1 = 0$, market 1 does not respond to a lagged disequilibrium error which reflects transitory movements away from the permanent component and CS_2 is zero.[8] Hence, CS_1 reflects how sensitive market 2 is relative to market 1 to lagged transitory shocks and vice versa (Yan and Zivot 2010).

Since IS measures are defined as the proportion of the efficient price innovation variance that can be attributed to specific prices they must include information from the covariance matrix. The reduced form VECM covariance matrix of the error terms $\boldsymbol{\varepsilon_t} = (\varepsilon_{1t}, \varepsilon_{2t})$, is

$$\boldsymbol{\Sigma} = \begin{bmatrix} \sigma_1^2 & \rho\sigma_1\sigma_2 \\ \rho\sigma_2\sigma_1 & \sigma_2^2 \end{bmatrix} = \begin{bmatrix} m_{11} & m_{12} \\ m_{21} & m_{22} \end{bmatrix} = \mathbf{M}, \tag{11}$$

where $\sigma_1^2(\sigma_2^2)$ is the variance of $\varepsilon_{1t}(\varepsilon_{2t})$ and ρ is the correlation. To establish the importance of the each series independently in price discovery, assume that $\boldsymbol{\Sigma}$ is diagonal. In this case, the IS measures are

$$IS_1 = \frac{(\gamma_1 m_{11})^2}{(\gamma_1 m_{11})^2 + (\gamma_2 m_{22})^2}, IS_2 = \frac{(\gamma_2 m_{22})^2}{(\gamma_1 m_{11})^2 + (\gamma_2 m_{22})^2}, \tag{12}$$

which shows the measures are the proportion of the total variance attributed to each price weighted by their respective CS measures. In effect, IS is the contribution of each price to the efficient innovation variance scaled or weighted by the response of each series to deviations or transitory movements from the equilibrium relationship.[9] A low (high) information share for a market implies a small (large) reaction to the arrival of new information about fundamental value. To implement the IS measures, researchers use the Cholesky factorization which will be identical to the measures presented when the correlation in the error terms is zero. In the presence of correlation, average measures are calculated based on different orderings of the VECM. Several studies have compared the CS and IS measures in practical applications. Studies by Baillie et al. (2002), Harris, McInish, and Wood (2002b), and

[8] This interpretation of CS links price discovery to weak exogeneity for the cointegrating parameters in a market (Zivot 2000).

[9] Baille et al. (2002) show that the vector of permanent component weights ($\boldsymbol{\gamma}$) in the Gonzalo-Granger PT decomposition and the vector of long-run impact coefficients ($\boldsymbol{\psi}$) that make up the efficient price innovation in Hasbrouck's IS framework are equal up to a scale factor. Hence, IS measures can also be defined in terms of elements of $\boldsymbol{\gamma}$.

Lehman (2002) indicate that CS and IS measures differ most when the error terms from the VECM are correlated. Hasbrouck (2003) also suggests that the results from price discovery metrics should be similar to those from impulse response functions because both measures are non-linear functions of the same parameters.

Recently, Yan and Zivot (2010) argue that the two methods alone cannot distinguish the price discovery dynamics between markets and show a method to use the CS and IS to disentangle the impact of permanent and transitory shocks.[10] They interpret CS as a measure of relative noise where the price with a lower CS is relatively more sensitive to transitory shocks. This interpretation is consistent with the Putnins' (2013) definition of CS as a measure of relative noise, and IS as a combination of relative noise and relative leadership in reflecting innovations in the fundamental value. In this situation, the Yan and Zivot provide an information leadership metric which combines both CS and IS to cancel out the relative noise and identify which series provides more information to the efficient price. Yan and Zivot's metric is

$$IL_1 = \left| \frac{IS_1}{IS_2} \frac{CS_2}{CS_1} \right|, IL_2 = \left| \frac{IS_2}{IS_1} \frac{CS_1}{CS_2} \right|. \tag{13}$$

Unlike the CS and IS, these measures are not shares because IL_1 and IL_2 do not add to one. IL_1 has the range $[0, \infty)$, and values of IL_1 above (below) one indicate that the price leads (does not lead) the process of incorporating new information into prices. To make the information leadership metric comparable to CS and IS, Putnins (2013) defines informational leadership shares (ILS)

$$ILS_1 = \frac{IL_1}{IL_1 + IL_2}, ILS_2 = \frac{IL_2}{IL_1 + IL_2}. \tag{14}$$

We use the above ILS metric to measure proportional contribution to price discovery between live cattle futures and cutout prices. The ILS_1 has the range $[0, 1]$, and values of ILS_1 above (below) .5 suggests that the first price leads (does not lead) the process of

[10] For further details see the work by Baillie et al. (2002), Harris, McInish, and Wood (2002a), Lehman (2002), Yan and Zivot (2010), Putnins (2013), and the original work by Hasbrouck (1995) and Gonzalo and Granger (1995).

incorporating new information (Putnins 2013).[11]

Empirical Results

Diagnostic Testing

We conduct unit root tests to identify the order of integration in futures and cutout prices. Augmented Dickey-Fuller (ADF) tests are performed with the number of lags chosen using the Schwarz Bayesian Information Criterion (SBIC). For the ADF test, where the dependent variable is differenced, we focus on the specification with a constant (Wang and Tomek 2007). ADF-GLS tests (Elliot, Rothenberg, and Stock 1996) are also used, which are ADF tests on GLS de-trended data. A specification with only a constant is used for the ADF-GLS test, and the Modified Akaike Information Criterion (MAIC) is used to identify lags for the tests. The null for ADF and ADF-GLS is that the series is non-stationary. The Kwiatkowski–Phillips–Schmidt–Shin (KPSS) tests which has a null hypothesis that the time series is stationary around a deterministic trend is also used. Finally, variants of the Zivot-Andrews (ZA) test which allows for one possible shift in the mean, trend, or both mean and trend are also used (Zivot and Andrews 2002). The ZA test has the null hypothesis of a unit root process with drift that excludes exogenous structural change. The alternative hypothesis is that a structural break may be present. The rejection of the null would imply rejection of a unit root without breaks. Both futures and cutout prices are tested for stationarity using the full sample of 508 observations spanning ten years.

Table 1 (panel A) presents the results of unit-root tests for contemporaneous Friday futures and cutout series. ADF, ADF-GLS, and KPSS tests on futures price series and cutout prices indicate a non-stationary process. The conclusions drawn from different specifications of the ZA test are mixed. The ZA test with a break in both intercept and trend rejects the null hypothesis that futures price in levels contains a unit root. However, the specifications

[11] The ILS measure has the advantage of not producing extreme values which occur in IL when both IS and CS approaches the value 1 (Putnins 2013).

with either a break in the intercept or trend indicate that futures price series has a unit root. In the case of cutout, ZA test specifications with an intercept or trend reject the null hypothesis of the unit root at the .05 level. On the other hand, ZA test allowing a break in both level and trend fails to reject the null hypothesis of unit root.[12] Figure 1 presents the nominal price series for contemporaneous futures and cutout price for the 1/9/2004-9/27/2013 period. Visual inspection of figure 1 reveals that the relationships between the futures and cutout markets may have varied over time with the futures more consistently above the cutout prices after 2011. Based on test results we conclude that both price series are non-stationary in levels and stationary in first differences. The results are consistent for Friday cutout prices and one day lagged futures prices (panel B).

Results of Cointegration Analysis

We focus on the cointegration relationship between futures and cutout prices for the full period and then test further for structural stability in the cointegration relationship. Enders (2008) suggests using a drift term outside the cointegrating relationship if the variables exhibit a decided tendency to increase or decrease. This allows the rank of π to be viewed as the number of cointegrating relationships after purging any linear trend in the data generating process. Hence, the maximum-likelihood estimation procedure by Johansen and Juselius (1990) is performed with an unrestricted intercept term, where a drift term appears in the level series (Luthkepohl 2006; Enders 2008). Based on SBIC, we use a specification with 3 lags for the test. The results are presented in table 2, panel A for the Friday futures and cutout prices, and panel B for Friday cutout prices and the Thursday futures prices. Both the trace and the λ-max tests indicate that there is one cointegrating vector at the .10, .05, and .01 levels of significance in both panels. The cointegration properties are also assessed using the Engle and Granger (1987) two-step procedure and the null of no cointegration is

[12]The break dates identified in the ZA test are not consistent across the different specifications. Nonetheless, the financial market crash in October 2008 is observed to have a large impact on the futures price series. In contrast, a period of rising beef demand towards the end of 2010 appears to have an impact on cutout prices.

rejected in panel A and panel B.[13] Next, we apply the Bierens and Martins (2010) procedure which tests the null of a time-invariant cointegrating vector. Since the futures and cutout prices appear to follow a unit root with drift process, the test is conducted under the drift case assumption. We fail to reject the null of a time-invariant cointegrating vector at the .05 significance level in both panels. These results are robust to several lags lengths (SBIC specified number of lags 3) and to Chebyshev polynomials up to 15 which was viewed as reasonable given the use of weekly observations.

Based on the cointegration test, a VECM is estimated for the full period using both samples imposing one cointegrating vector. While the SBIC chose 3 lags, an additional lag is included to control autocorrelation.[14] Results of weak exogenetiy tests are presented in table 3. We fail to reject weak exogeneity of futures prices at the .01 level in panel A., indicating that the current futures price does not make short-run adjustments to the long-run disequilibrium. In contrast, weak exogeneity of the cutout market is rejected at the .01 level, reflecting the adjustments to the disequilibrium. The results are consistent for current cutout prices and one-day lagged futures prices in panel B. The speed-of-adjustment parameter for cutout is substantially higher than that for futures prices in both samples reflecting the responsiveness of cutout prices to last period's equilibrium error.[15]

Results of Innovation Accounting

Forecast error variance decompositions and impulse response functions are employed to reveal the magnitude of short-run linkages between markets. The Cholesky factorization is

[13] These results are not reported and are available on request.

[14] There is no serial correlation in the residuals up to 6 lags for both models at the .05 level. While mild ARCH effects are present in the residuals of the equivalent level VAR with the cointegration restriction imposed, Gonzalo (1994) demonstrates that cointegration conclusions based on Johansen's maximum-likelihood estimation procedure are robust.

[15] An Autoregressive Distributed Lag (ARDL) model is also estimated imposing the weak exogeneity. The model with SBIC selected 2 lags has an $R^2 = .39$ indicating a reasonable statistical relationship between futures and cutout prices. The residuals of the model are homoscedastic with no autocorrelation up to 10 lags. The speed of adjustment parameter for the cutout price estimated in the ARDL model is .08 which similar to our estimates from the VECM. The results for Friday cutout prices and one-day prior futures price are similar with a marginally higher $R^2 = .41$ and some residual heteroskedasticity.

employed to assign the causal ordering recursively for futures and cutout. The innovation vector from an equivalent levels VAR model can be written as $\mathbf{A}\boldsymbol{\upsilon_t} = \boldsymbol{\epsilon_t}$ where $\mathbf{A}$ is a (2×2) matrix and $\boldsymbol{\upsilon}_t$ is a vector of orthogonal shocks. Ordering the futures price first, the $\mathbf{A}$ matrix gives the following representation on innovations in contemporaneous time

$$\begin{bmatrix} 1.0 & 0.0 \\ a_{21} & 1.0 \end{bmatrix} \times \begin{bmatrix} \upsilon_{f,t} \\ \upsilon_{b,t} \end{bmatrix} = \begin{bmatrix} \epsilon_{f,t} \\ \epsilon_{c,t} \end{bmatrix}, \tag{15}$$

where $\epsilon_{i,t}$ terms are the observed innovations from VECM, $\upsilon_{i,t}$ are the orthogonal innovations from each market, and i = futures (f) and cutout (c) respectively. Enders (2008) and Luthkepohl (2006) note that the ordering of the variables can influence forecast error variance decompositions and impulse response functions particularly if the correlations between the innovations exceed $|.20|$. We find that the correlation between the VECM residuals is .26 for the first sample and .35 for the second sample which are significant at the .01 level. Hence, we also assess the sensitivity of our results ordering cutout price first which restricts the contemporaneous effect of futures price on cutout prices and allow the cutout price to be "causally prior" to the futures price.

The 15-week forecast error variance decompositions for the full sample are reported in table 4. The forecast error variance decompositions identify the proportion of the movement in a particular sequence due to its "own" shocks versus shocks to other variables (Enders 2008). Since it is common for a variable to explain almost all of its own forecast error variance at short horizons and smaller proportions at longer horizons, the real test in terms of market dominance lies in the longer horizons. The forecast error variance in each price series (listed vertically) is decomposed into proportions due to shock in futures and cutout and reported for a 15-week horizon. For the contemporaneous futures and cutout price sample (panel A), futures price contributes 100% of its own forecast error variance at the first-week horizon and 99.7% of its own forecast error variance at the longer 15-week horizon. This implies that the futures market is highly exogenous i.e., the futures price evolves independently of the forecast error shocks from the cutout market. In the case of the cutout market, the forecast errors in the first-week horizon (93.4%) can be largely attributed to its own innovations in the first week. However, as we move towards the 15-week horizon, the futures price dominates by explaining 60% of the variance in the cutout market, with cutout explaining only 40% of its own variance. The results for the Friday cutout and one-day prior futures price sample in panel B are consistent with panel A. Alternate ordering with cutout prices ordered first marginally increases the contribution of cutout in its own variance decomposition with the futures price dominating in the longer horizon for both panels.

The impulse response functions are plotted with 95% bootstrapped confidence intervals (2000 runs) for the 15-week horizon and presented for the contemporaneous futures and cutout price (figure 2). Impulse response functions trace the time path of the various shocks on the prices included in the VAR system. The responses to a one standard deviation shock in the futures market are presented in the first row, followed by responses to shocks from the cutout market.[16] The responses obtained are consistent with the results observed from forecast error variance decompositions. The futures and cutout markets respond significantly to the shocks from the futures markets i.e., a one standard deviation shock in the futures market produces responses that are significantly different from zero at the .05 level in the cutout market at all horizons. In contrast, the shock in the cutout market does not influence the futures market at any horizon. The results are consistent for the Friday cutout price and one-day prior futures price series and are not reported. Further assessment ordering the cutout price first does not appear to cause any difference in the impulse response functions. Overall, the results of innovation accounting remain robust and support the dominant role of futures prices.

Price Discovery Measures

Examination of equations (10), (11), (13), and (14) reveal the importance of weak exogeneity

[16] The impulse responses may also be interpreted in percentage terms after scaling with the corresponding VECM residual standard deviation (SD) for futures (.0229) and cutout price (.0195). For example, a 1-SD shock in the futures price innovation causes a response of .015 in cutout price which is equivalent to a 1% shock in futures causing a response of .66% in the cutout price.

in the price discovery metrics. When a price series is weakly exogenous, it's corresponding α in the VECM is equal to zero which using (10) translates into a CS equal to zero ($\gamma = 0$) for the other price series. Because these gammas are embedded in the IS measures, the complete dominance of a series reflected in CS is transmitted to the IS measures, and because one of the IS measures will be zero, interpretation of (13) and (14) becomes problematic. In our situation, the weak exogeneity test statistically attributes all weight in price discovery to the cattle futures market using either the Friday or Thursday data.

Despite these findings, we use the estimated speed of adjustment coefficients to calculate the price discovery measures, generating a maximum cutout price effect in the price discovery process. Based on the estimated coefficients, the CS values computed are 90.5% for the Friday futures and 9.5% for Friday cutout. For Thursday futures and Friday cutout, the CS values are 92.2% and 7.8% respectively. In both cases, CS identifies the dominance of the futures price in the price discovery process. The IS for Friday futures and Friday cutout is 94% and 6% respectively. For Thursday futures price and Friday cutout price, futures continue to dominate contributing 91.5% of price discovery and cutout the remaining 8.5%.[17] The higher values of IS indicates that the futures price is quicker in incorporating new information compared to cutout price. The close approximation of proportional contributions by both CS and IS measures point to the relatively similar variances and rather low contemporaneous correlation in the reduced form residuals. As Hasbrouck (2003) suggests, the results from price discovery metrics should be similar to those from the impulse response analysis. In both cases, the change in the cutout price does not have an effect on the futures price.

Based on the Yan and Zivot (2010) and Putnins (2013) approaches, we use these CS and IS measures to derive the ILS measures. The computed ILS measures confirm the

[17] The IS values reported are the averages from alternate ordering of futures and cutout prices. For both contemporaneous and one-day prior futures price relationship, the futures price is dominant contributing more than 99.3% to price discovery with the futures price ordered first and more than 83.5% with the cutout price ordered first.

dominance of futures price in reflecting new information.[18] The ILS for the Friday futures price in the cutout price is 73.1%, with the cutout price contributing only 26.9% to the price discovery. A smaller ILS value for futures compared to CS and IS is attributed to the elimination of the relatively higher transitory noise from the cutout price. When the sample of Friday cutout prices and Thursday futures price is compared, the contribution to price discovery from cutout price increases to 54.9% which is marginally higher than the 45.1% contribution of futures price. While this change in the importance of futures prices in price discovery might be attributed to a loss of information by using the Thursday rather than the Friday price as well as the elimination of noise from the Friday cutout price, the magnitude of the reduction seems large. Recalculation of highly non-linear IL on which the ILS is based, and examination of their components also suggest that IL is highly sensitive near the bounds to even small changes in CS and IS values. In short, the new price discovery measures that allow for a maximum advantage to cutout price in the pricing process still identify the dominance of the current futures price, and about equal weighting for the lagged one-day futures price. Nevertheless, recall that the results of the weak exogeneity tests attribute all the weight in the price process to the futures price.[19]

Conclusions

We investigate the relationship between the CME live cattle futures price and the cutout price for the 2004-2013 post-LMRA period to identify the price market participants should watch for timely and reliable cattle price information. Our findings indicate that the live cattle futures price is considerably more informative than the cutout price. The two series are cointegrated by one long-run time-invariant vector. The dominance of the futures price in the price relationship is supported by weak exogeneity tests which indicate that the futures price does not adjust to the discrepancy from long-run equilibrium. The importance of the

[18] We report the ILS values as they are easier to interpret. The IL values are available on request.

[19] As the market is known to be under transition, robustness checks were done dividing the data into two equal samples. In addition, the analysis was also performed employing a contract roll dummy. These results are consistent with previous findings and are available on request.

futures price is strengthened by innovation accounting using the structural VECM. Forecast error variance decompositions reveal that the futures market is not affected by shocks in the cutout market and plays a dominant role in cutout decompositions at distant horizons. Shocks to the futures price innovations also exhibit a relatively strong and lasting effect in the cutout market. In contrast, shocks to the cutout innovations do not influence futures prices, and cutout forecast errors contribute practically nothing to the error variance in the futures market. Based on the weak exogeneity tests, newly-developed price discovery metrics attributes the dominant role to the futures market. Even when these metrics are interpreted to allow for a maximum cutout effect in the pricing process, the dominance of the current futures price emerges and indicates about equal weighting for the lagged one-day futures price. Our results indicate that the futures price adjusts to new fundamental information quicker than cutout price.

While we do find that the cutout price enters into the long-run cointegrating relationship, its limited importance in short-run dynamics was somewhat unexpected. Since the cutout prices are important in calculating processing margins, this disconnect in the short-run may be related to episodic non-competitive changes in processing margins documented by Cai, Stiegert, and Koontz (2011a, 2011b). From a different perspective, the limited importance also may be due to shortcomings in reported boxed beef cutout values. Wholesale cutout values do not include prices for the growing and variable exported beef products that on average are higher value cuts. For instance, branded products such as the Certified Angus Beef brand which uses the upper 2/3 Choice make up a significant volume of beef trade and is not included in Choice boxed beef cutouts. Such reporting issues may cause the wholesale values to underestimate the true, but more volatile value of wholesale meat. It also remains uncertain whether the thinness of the reported wholesale cutout prices resulting from the shift to contracting at retailer/wholesaler level (Koontz and Ward 2011) have contributed to the limited representativeness of reported cutout prices.

In sum, the futures price has a strong predictive influence on the cutout price and appears to assimilate fed cattle price information quicker than both current and one-day ahead cutout prices. From a weekly pricing perspective, the cutout price computed from cutout values do not inform the live cattle futures price. The implication for a producer marketing cattle in the coming week is that the futures price provides the most relevant and timely source of fundamental information. While the cutout price appears to have limited value in short-term price discovery, the extensive information available in boxed beef reports is valuable in understanding general patterns in beef demand and supply along the marketing chain. For instance, the spread between the Choice and Select cutout informs the market on the relative supply of each grade informing about packer demand for cattle with specific characteristics. Information on changes in volume and price may also reflect market dynamics including changing buyer preference, price resistance along the marketing chain, as well as glut or products backing-up in distribution pipeline (USDA-AMS). Regardless, the beef markets have evolved over time and there may be a need to evaluate the way in which boxed beef cutout values are developed to make prices more timely and informative.

References

Azzam, A. 2003. "Market Transparency and Market Structure: The Livestock Mandatory Reporting Act of 1999." *American Journal of Agricultural Economics* 85(2): 387-395.

Baillie, R.T., G.G. Booth, Y. Tse, and T. Zabotina 2002. "Price Discovery and Common Factor Models." *Journal of Financial Markets* 5(3): 309-321.

Becker, G.S. 2006. "Livestock Price Reporting: Background." CRS Report for Congress 2006. Congressional Research Service, Washington, D.C.

Beveridge, S. and C.R. Nelson. 1981. "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to Measurement of the 'Business Cycle'." *Journal of Monetary Economics* 7(2): 151-174.

Bierens, H.J. and L.F. Martins. 2010. "Time-Varying Cointegration." *Econometric Theory* 26(05): 1453-149.

Booth, G.G., R.W. So, and Y. Tse. 1999. "Price Discovery in the German Equity Index Derivatives Markets." *Journal of Futures Markets* 19(6): 619-643.

Boyd, M. and B.W. Brorsen. 1985. "Dynamics of Weekly Prices in the United States Beef and Pork Marketing Channels." *Canadian Journal of Agricultural Economics* 33(3): 331-342.

Boyer, C.N. and B.W. Brorsen. 2013. "Changes in Beef Packers' Market Power after the Livestock Mandatory Price Reporting Act: An Agent-Based Auction." *American Journal of Agricultural Economics* 95(4): 859-876.

Cai, X., K.W. Stiegert, and S.R. Koontz. 2011a. "Regime Switching and Oligopsony Power: The Case of U.S. Beef Processing." *Agricultural Economics* 42(1): 99-109.

____. 2011b. "Oligopsony Fed Cattle Pricing: Did Mandatory Price Reporting Increase Meatpacker Market Power?" *Applied Economic Perspectives and Policy* 33(4): 606-622.

Chu, Q. C., W.L.G. Hsieh, and Y. Tse. 1999. "Price Discovery on the S&P 500 Index

Markets: An Analysis of Spot Index, Index Futures, and SPDRs." *International Review of Financial Analysis* 8(1): 21-34.

CME Rulebook, Chapter 101 Live Cattle Futures. [Accessed on 05/27/2013.] http://www.cmegroup.com/rulebook/CME/II/100/101/101.pdf

Dolado, J.J. and H. Lutkepohl. 1996. "Making Wald Tests Work for Cointegrated VAR Systems." *Econometric Reviews* 15(4): 369-386.

Elliott, G., T.J. Rothenberg, and J.H. Stock. 1996. "Efficient Tests for an Autoregressive Unit Root." *Econometrica* 64(4): 813-836.

Enders, W. 2008. "Applied Econometric Time Series." *John Wiley & Sons.*

Engle R.F. and C.W.J. Granger 1987. "Co-integration and Error Correction: Representation, Estimation, and Testing." *Econometrica* 55(2): 251-276.

Garbade, K.D. and W.L. Silber. 1983. "Price Movements and Price Discovery in Futures and Cash Markets." *Review of Economics and Statistics* 65(2): 289-297.

Garcia, P., R.M. Leuthold, T.R. Fortenbery, and G.F. Sarassoro. 1988. "Pricing Efficiency in the Live Cattle Futures Market: Further Interpretation and Measurement." *American Journal of Agricultural Economics* 70(1): 162-169.

Grunewald, S., T.C. Schroeder, and C.E. Ward. 2004. "Cattle Feeder Perceptions of Livestock Mandatory Price Reporting." *Applied Economic Perspectives and Policy* 26(4): 521-538.

Goodwin, B.K. and M.T. Holt. 1999. "Price Transmission and Asymmetric Adjustment in the U.S. Beef Sector." *American Journal of Agricultural Economics* 81(3): 630-637.

Gonzalo, J. 1994. "Five Alternative Methods of Estimating Long-Run Equilibrium Relationships." *Journal of Econometrics* 60(1-2): 203-233.

Gonzalo, J. and C.W.J. Granger. 1995. "Estimation of Common Long-Memory Components in Cointegrated Systems." *Journal of Business and Economic Statistics* 13(1): 27-36.

Hahn, W.F. 1990. "Price Transmission Asymmetry in Pork and Beef Markets." *Journal of Agricultural Economics Research* 42(4): 21-30.

Hamilton, J.D. 1994. "Time Series Analysis." *Princeton University Press, Princeton, NJ.*

Hansen, B.E. 1992. "Testing for Parameter Instability in Linear Models." *Journal of Policy Modeling* 14(4): 517-533.

Harris, F.H.D, T.H. McInish, and R.A. Wood. 2002a. "Security Price Adjustment across Exchanges: An Investigation of Common Factor Components for Dow Stocks." *Journal of Financial Markets* 5(3): 277-308.

____. 2002b. "Common Factor Components versus Information Shares: A Reply." *Journal of Financial Markets* 5(3): 341-348.

Hasbrouck, J. 1995. "One Security, Many Markets: Determining the Contributions to Price Discovery." *Journal of Finance* 50(5): 1175-1199.

____. 2003. "Intraday Price Formation in U.S. Equity Index Markets." *The Journal of Finance* 58(6): 2375-2400.

Ji, I.B. and C. Chung. 2012. "Causality between Captive Supplies and Cash Market Prices in the U.S. Cattle Procurement Market." *Agricultural and Resource Economics Review* 41(3): 340-350.

Johansen, S. 1991. "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models." *Econometrica* 59(6): 1551-1580.

____. 1992. "Determination of Cointegration Rank in the Presence of a Linear Trend." *Oxford Bulletin of Economics and Statistics* 54(3): 383-397.

Johansen, S. and K. Juselius. 1990. "Maximum Likelihood Estimation and Inference on Cointegration–with Applications to the Demand for Money." *Oxford Bulletin of Economics and Statistics* 52(2): 169-201.

____. 1992. "Testing Structural Hypothesis in a Multivariate Cointegration Analysis of the

PPP and the UIP for UK." *Journal of Econometrics* 53(1): 211-244.

Koontz, S.R., P. Garcia, and M.A. Hudson 1990. "Dominant-Satellite Relationships between Live Cattle Cash and Futures Markets." *Journal of Futures Markets* 10(2): 123-136.

Koontz, S.R. 2007. "Impacts of Mandatory Price Reporting on the Relationship between Fed Cattle Prices and the USDA Boxed Beef Cutout Value." Working Paper, Department of Agricultural and Resource Economics, Colorado State University.

Koontz, S.R. and C.E. Ward. 2011. "Livestock Mandatory Price Reporting: A Literature Review and Synthesis of Related Market Information Research." *Journal of Agricultural and Food Industrial Organization* 9(1): 1-33.

Lee, Y., C.E. Ward, and B.W. Brorsen. 2012. "Procurement Price Relationships for Fed Cattle and Hogs: Importance of the Cash Market in Price Discovery." *Agribusiness* 28(2): 135-147.

Lehmann, B.N. 2002. "Some Desiderata for the Measurement of Price Discovery across Markets." *Journal of Financial Markets* 5(3): 259-276.

Lutkepohl, H. 2006. "New Introduction to Multiple Time Series Analysis." *Springer New York.*

McKenzie, A.M. and M.T. Holt. 2002. "Market Efficiency in Agricultural Futures Markets." *Applied Economics* 34(12): 1519-1532.

Marsh, J.M. and G.W. Brester. 1989. "Intertemporal Price Adjustment in the Beef Market: A Reduced Form Analysis of Weekly Data." *Western Journal of Agricultural Economics* 14(2): 235-245.

Martinez, V., P. Gupta, Y. Tse, and J. Kittiakarasakun. 2011. "Electronic Versus Open Outcry Trading in Agricultural Commodities Futures Markets." *Review of Financial Economics* 20(1): 28-36.

Njoroge, K. 2003. "Information Pooling and Collusion: Implications for the Livestock Mandatory Reporting Act." *Journal of Agricultural & Food Industrial Organization* 1(1): 1-15.

Njoroge, K., A. Yiannaka, K. Giannakas, and A.M. Azzam. 2007. "Market and Welfare Effects of the U.S. Livestock Mandatory Reporting Act." *Southern Economic Journal* 74(1): 290-311.

Oellerman, C.O, B.W. Brorsen, and P.L. Farris 1989. "Price Discovery for Feeder Cattle." *Journal of Futures Markets* 9(2): 113-121.

Oellerman, C.O. and P.L Farris. 1985. "Futures or Cash: Which Market Leads Live Beef Cattle Prices?" *Journal of Futures Markets* 5(4): 529-538.

Park, M., Y. Jin, and A.H. Love. 2011. "Dynamic and Contemporaneous Causality in a Supply Chain: An Application of the U.S. Beef Industry." *Applied Economics* 43(30): 4785-4801.

Perry, J., J. MacDonald, K. Nelson, W. Hahn, C. Arnade, and G. Plato. 2005. "Did Mandatory Requirement Aid the Market?" Economic Outlook Report No. LDP-M-135-01, Economic Research Service, U.S. Department of Agriculture, Washington, D.C.

Phillips, P.C.B. 1998. "Impulse Response and Forecast Error Variance Asymptotics in Non-stationary VARs." *Journal of Econometrics* 83(1): 21-56

Putnins, T.J. 2013. "What do Price Discovery Metrics Really Measure?" *Journal of Empirical Finance* 23: 68-83.

Schroeder, T.C. and B.K. Goodwin. 1991. "Price Discovery and Cointegration for Live Hogs." *Journal of Futures Markets* 11(6): 685-696.

Schroeder, T.C. and J. Mintert. 2000. "Alternatives to Cash Prices in Fed Cattle Price Discovery." Conference Paper, Cattlemen's Day, Kansas State University: 93-94.

Schroeder, T.C. and M.L. Hayenga. 1987. "Short-Term Vertical Market Price Interrelationships for Beef and Pork." *North Central Journal of Agricultural Economics* 9(2): 171-180.

Sims, C. 1986. "Are Forecasting Models Usable for Policy Analysis." *Quarterly Review of*

Federal Reserve Bank of Minneapolis. 10(1): 2-16.

Stock, J.H. and M.W. Watson. 1988. "Testing for Common Trends." *Journal of the American Statistical Association* 83(404): 1097-1107.

Swanson, N. and C.W.J. Granger. 1997. "Impulse Response Functions Based on a Causal Approach to Residual Orthogonalization in Vector Autoregression." *Journal of American Statistical Association* 92(437): 357-367.

Toda, H.Y. and T. Yamamoto. 1995. "Statistical Inference in Vector Autoregressions with Possibly Integrated Processes." *Journal of Econometrics* 66(1): 225-250.

Wachenheim, C.J. and E.A. DeVuyst. 2001. "Strategic Response to Mandatory Reporting Legislation in the U.S. Livestock and Meat Industries: Are Collusive Opportunities Enhanced?" *Agribusiness* 17(2): 177-195.

Wang, D. and W.G. Tomek. 2007. "Commodity Prices and Unit Root Tests." *American Journal of Agricultural Economics* 89(4): 873-889.

Ward, C.E. 1981. "Short Period Pricing Model for Fed Cattle and Impacts of Wholesale Carcass Beef and Live Cattle Futures Market Prices." *Southern Journal of Agricultural Economics* 13(1): 125-132.

____. 2006. "An Assessment of the Livestock Mandatory Reporting Act." *Proceedings of the NCCC-134 Conference on Applied Commodity Price Analysis, Forecasting, and Market Risk Management.*

____. 2010. "Assessing Competition in U.S. Beef Packing Industry." *Choices* 25(2): 1-14.

Ward, C.E., T.C. Schroeder, J. Mintert, and D.S. Peel. 1997. Fed Cattle Price Discovery Issues and Projections. Oklahoma Cooperative Extension Service, Oklahoma State University.

Yan, B. and E. Zivot. 2010. "A Structural Analysis of Price Discovery Measures." *Journal of Financial Markets* 13(1): 1-19.

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

Zivot, E. 2000. "The Power of Single Equation Tests for Cointegration When the Cointegrating Vector is Prespecified." *Econometric Theory* 16(03): 407-439.

Zivot, E. and D.W.K. Andrews. 2002. "Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis." *Journal of Business and Economic Statistics* 20(1): 25-44.

Table 1. Unit Root Tests: Weekly U.S. Cattle Prices

Market	Test	Level		First Difference		Level		First Difference	
		Panel A.				Panel B.			
Futures	ADF (C)	-1.59		-24.28	***	-1.62		-23.73	***
	ADF GLS (C)	0.03		-1.24		0.02		-1.17	
	PP(CT)	-2.79		-24.31	***	-2.82		-23.79	***
	KPSS (C)	2.21	***	0.04		2.22	***	0.04	
	ZA (C)	-4.65		-		-4.65		-	
	ZA (T)	-4.14		-		-4.15		-	
	ZA (CT)	-5.30	**	-		-5.26	**	-	
Cutout	ADF (C)	-1.67		-21.29	***	-1.65		-21.51	***
	ADF GLS (C)	-0.38		-0.07		-0.15		-0.07	
	PP(CT)	-3.78	**	-17.92	***	-3.77	**	-17.81	***
	KPSS (C)	2.22	***	0.03		2.22	***	0.03	
	ZA (C)	-5.12	**	-		-5.06	**	-	
	ZA (T)	-4.48	**	-		-4.43	**	-	
	ZA (CT)	-5.00		-		-4.95		-	

Notes: Data: 1/9/2004-9/27/2013. Panel A: Friday futures and cutout price. Panel B: Thursday futures price and Friday cutout price. Futures denotes CME live cattle futures price. Cutout denotes the cutout values multiplied by 63% to reflect the average hot yield. C denotes a specification with a constant. T denotes a specification with a trend. CT denotes a specification with both constant and trend. The τ-stat is the test statistic for the ADF test. Lag lengths for the ADF test are based on SBIC. ADF-GLS (C) is specified with an intercept. The test statistic for ADF-GLS is t-stat. Lag lengths for ADF-GLS are based on the MAIC. ZA test is specified with one-time break in both intercept and trend. Lags for ZA test are based on SBIC. For the ZA test, the test statistic is t-stat. ***Significant at $\alpha = 0.01$, **Significant at $\alpha = 0.05$, *Significant at $\alpha = 0.10$.

Table 2. Johansen Test for Cointegration

Null	Alternate			Critical Value		
Hypothesis	Hypothesis	Panel A.	Panel B.	90%	95%	99%
λ_{trace} *tests*		λ_{trace} *value*	λ_{trace} *value*			
r=0	r>0	30.18	31.5	15.66	17.95	23.52
r<=1	r>1	1.4	1.22	6.5	8.18	11.65
λ_{max} *tests*		λ_{max} *value*	λ_{max} *value*			
r=0	r=1	28.78	30.28	12.91	14.9	19.19
r=1	r=2	1.4	1.22	6.5	8.18	11.65

Notes: Data: 1/9/2004-9/27/2013. Panel A: Friday futures and cutout price. Panel B: Thursday futures price and Friday cutout price. Futures denotes CME live cattle futures price. Cutout denotes the boxed beef cutout values multiplied by 63% to reflect the average hot yield. LR test of the null hypothesis that there are at most "r" cointegrated vectors against the alternative that there are "k" cointegrated vectors (λ_{trace}) and, that there "r+1" cointegrated vectors (λ_{max}). The model is specified with an unrestricted constant term to account for possible trend (drift) in level series (Enders 2008).

Table 3. Weak Exogeneity Test

Panel A.					Panel B.				
ΔF (t)		ΔB (t)			ΔF (t)		ΔB (t)		
Coefficient (α)	-0.008	Coefficient (α)	0.081	***	Coefficient (α)	-0.007	Coefficient (α)	0.081	***

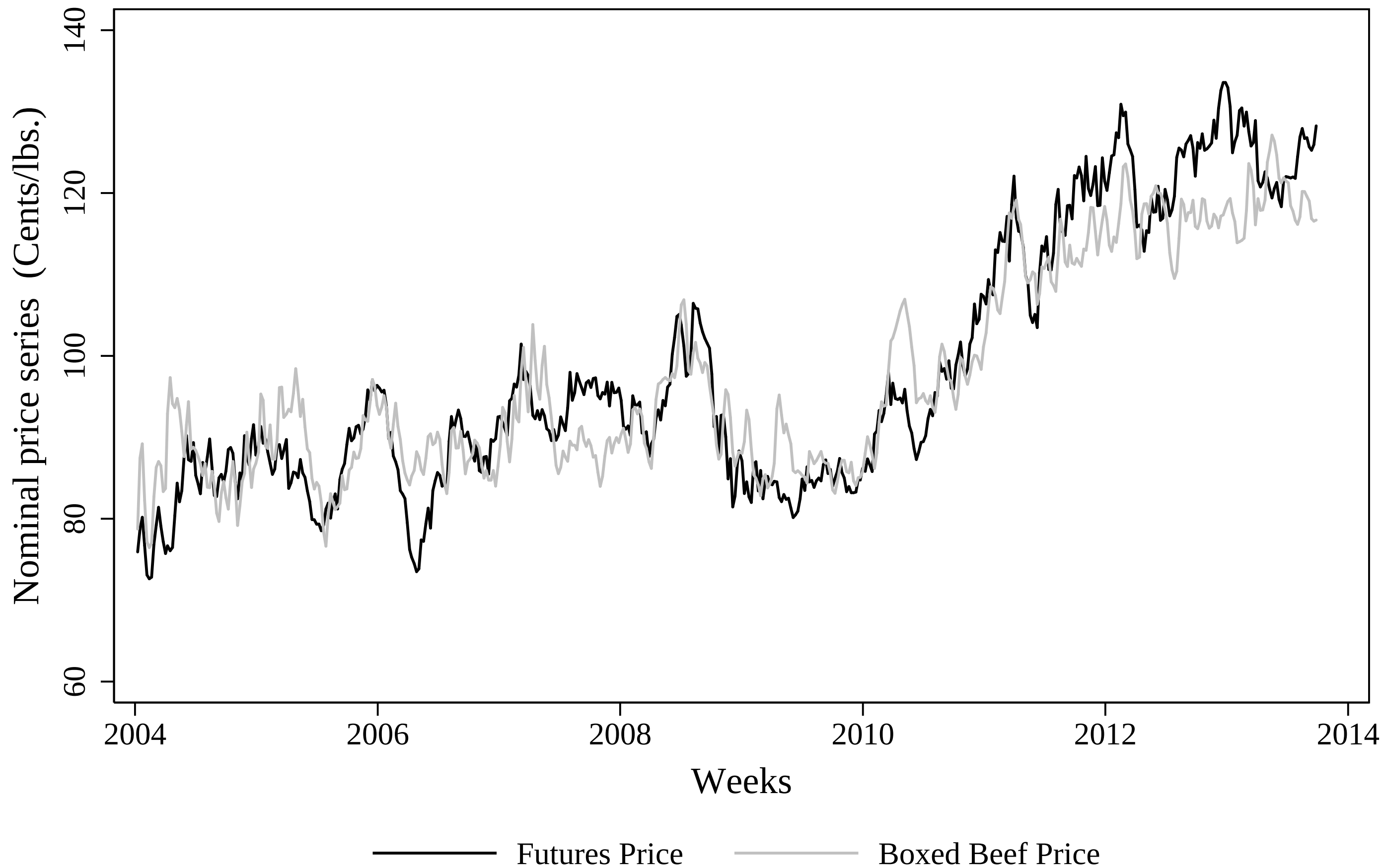
Notes: Data: 1/9/2004-9/27/2013. Panel A: Friday futures and cutout price. Panel B: Thursday futures price and Friday cutout price. Futures (F) denotes CME live cattle futures price. Cutout (C) denotes the boxed beef cutout values multiplied by 63% to reflect the average hot yield. The SBIC chosen lag is 3 for models in panels A and B. An additional lag is included to control autocorrelation. The test statistic is the Z statistic. ***Significant at $\alpha = 0.01$, **Significant at $\alpha = 0.05$, *Significant at $\alpha = 0.10$.

Table 4. Forecast Error Variance Decompositions From Level VAR

	Panel A			Panel B		
	Weeks	Futures	Cutout	Weeks	Futures	Cutout
Futures	1	100.00	0.00	1	100.00	0.00
	2	99.19	0.81	2	99.80	0.20
	5	99.23	0.77	5	99.32	0.68
	10	99.57	0.43	10	99.61	0.39
	15	99.68	0.32	15	99.72	0.28
Cutout	1	6.54	93.46	1	11.97	88.03
	2	20.00	80.00	2	24.09	75.91
	5	34.78	65.22	5	35.00	65.00
	10	49.47	50.53	10	49.60	50.40
	15	60.14	39.86	15	60.22	39.78

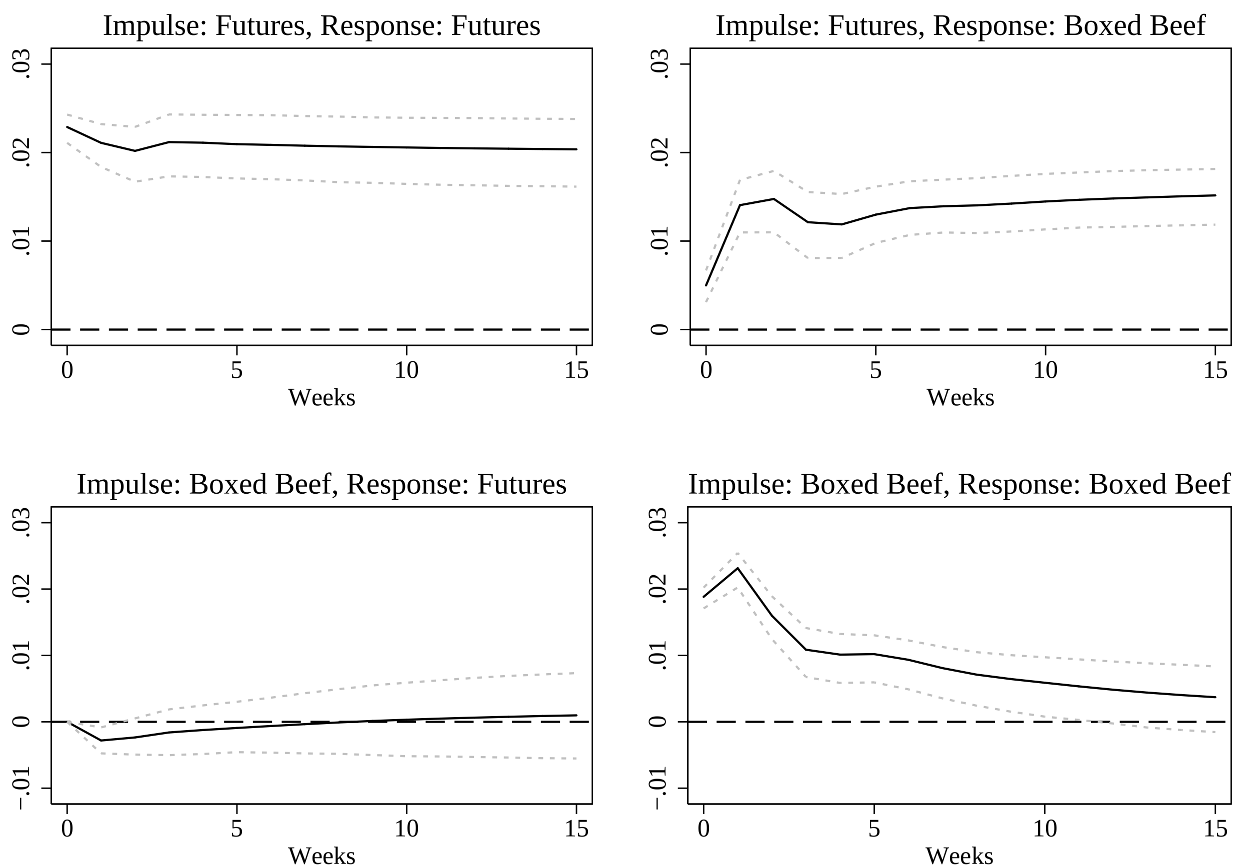
Notes: Data: 1/9/2004-9/27/2013. Panel A: Friday futures and cutout price. Panel B: Thursday futures price and Friday cutout price. Futures denotes CME live cattle futures price. Cutout denotes the boxed beef cutout values multiplied by 63% to reflect the average hot yield. The forecast error variance decompositions in each row sum to 100.

Figure 1. U.S. Fed Cattle Prices, January 2004-September 2013



Notes: Data: 1/9/2004-9/27/2013. Futures price denotes CME live cattle futures price. Cutout price denotes the boxed beef cutout values multiplied by 63% to reflect the average hot yield.

Figure 2. Impulse Response to One SD Innovations



Notes: Data: 1/9/2004-9/27/2013. Impulse response with 95% bootstrapped confidence intervals from level VAR. SD denotes standard deviation. Impulse response functions are reported for Friday futures and cutout prices.