The World's Largest Open Access Agricultural & Applied Economics Digital Library

This document is discoverable and free to researchers across the globe due to the work of AgEcon Search.

Help ensure our sustainability.

Give to AgEcon Search

AgEcon Search
http://ageconsearch.umn.edu
aesearch@umn.edu

*Papers downloaded from **AgEcon Search** may be used for non-commercial purposes and personal study only. No other use, including posting to another Internet site, is permitted without permission from the copyright owner (not AgEcon Search), or as allowed under the provisions of Fair Use, U.S. Copyright Act, Title 17 U.S.C.*

Food System Research Group

Working Paper Series

FSWP2010-2

REGIME SWITCHING AND OLIGOPSONY POWER:

THE CASE OF U.S. BEEF PROCESSING

Xiaowei Cai
California Polytechnic State University

Kyle W. Stiegert
University of Wisconsin-Madison

Steven R. Coonz
Colorado State University

In this article, we estimate a Green and Porter (GP 1984) trigger pricing model to analyze weekly marketing margin data for the U.S. beef packing industry from 1992-2000. Oligopsonists are hypothesized to follow a discontinuous pricing strategy in equilibrium, and we focus on shocks in the normal throughput of supply as a potential catalyst for regime switching between cooperative and noncooperative phases. We use an algorithm developed by Bellone (2005) that relies on Hamilton's (1989) multivariate first-order Markov process to test for the cooperative/noncooperative switching behavior. We find strong evidence that links switching conduct by packers to disruptions in coordinating the derived demands for processed beef with the supply of live cattle. Once switched, cooperative regimes lasted an average of 21 weeks, while noncooperative regimes averaged 33 weeks. The average marketing margin for processed beef was 66% higher in the cooperative regimes compared to the noncooperative regimes. This implies an annual average increase in profits of 491 million dollars to the beef packing industry.

JEL code: L11, L13, D43

Keywords: Margin, Beef Packing, Fed Cattle Prices, Markov Regime Switching

Food System Research Group
University of Wisconsin-Madison
http://www.aae.wisc.edu/fsrg/

January 2010

Authors emails:

cai@calpoly.edu
kwstiegert@wisc.edu
stephen.koontz@colostate.edu

This research was supported by a USDA grant to the Food System Research Group, University of Wisconsin, Madison.

FSRG *Working Papers* are published with the intention of provoking discussion and helping drive future work. The papers are formatted to conform with other working papers but are not formally edited.

All views, interpretations, recommendations, and conclusions expressed are those of the author(s) and not necessarily those of the supporting or cooperating institutions.

Copyright © by the author(s). All rights reserved.

Readers may make verbatim copies of this document for noncommercial purposes by any means, provided that this copyright notice appears on all such copies

1 Introduction

Trends toward increased concentration in beef packing from the 1970s through the 1990s raised major public policy concerns about the potential for beef packers to use monopsony power to lower fed cattle prices. In these decades, average plant size grew five-fold, the number of plants declined by over 70%,[1] and industry concentration, as measured by the Herfindahl-Hirschmann Index (HHI) for steer and heifer slaughter increased steadily. By 1999, the four largest beef packers accounted for 80% of the national beef slaughter, as opposed to 36% in 1980 (GIPSA 2002). In 1992 and for the first time, the national HHI index for steer and heifer slaughter rose above the threshold of 1800 defined by the U.S. Department of Justice as a highly concentrated industry (Whinston 2007).

Many studies have evaluated the potential for buyer power in U.S. beef packing. Azzam and Anderson (1996), Sexton (2000) and Ward (2002) provide excellent reviews of the literature through 2001. All three summaries suggest that the evidence supporting buyer power distortions is not conclusive. However, they are also critical of the methods employed in many of these studies. In general, there was concern about the structural rigidities implicit in structure-conduct-performance analysis and conjectural variation methods. Ward (2002) points toward a greater need to focus on the dynamic interactions among firms as a way to more flexibly identify potential price impacts from buyer power.

A failure to understand how market power is manifested spatially or over time would make it difficult to identify and measure with much precision. And this may be one factor driving the supposed inconclusive results from past work. At least three past studies have evaluated the beef packing industry in ways that capture dynamic patterns in buyer power behavior. Koontz, Garcia and Hudson (KGH 1993) used a noncooperative game model to evaluate breakdowns in oligopsony power using daily price movements. Evidence that cooperative regimes broke down and re-emerged over time supports strongly the point made by Ward (2002). Koontz and Garcia (KG 1997) later extended the single market model in KGH (1993) to multiple markets and found that low prices were paid in all relevant markets

[1] The average number of slaughter per plant increased from 32,383 in 1972 to 163,071 heads in 1998 (Ward and Schroeder 2002). The number of beef slaughter facilities declined from over 600 in 1980 to about 170 in 1999 (Barkema, Drabenstott and Novack 2001).

in cooperative phases, while high prices were paid in the noncooperative phases. Azzam and Park (1993) adopted traditional Bresnahan's procedure to test for switching market conduct in the beef slaughter industry. They found the evidence of market power by identifying the starting and ending points for the two distinct regimes of noncooperative and cooperative conduct. Other studies offer findings suggestive of regime switching.[2]

In the present paper, we analyze the beef procurement market using a Green and Porter (1984, hereafter GP) regime switching framework. Our model and analysis is in the tradition of KGH (1993) but differs from their work in three important ways. First, we use data from the 1992 period forward when national concentration levels were above the HHI=1800 DOJ threshold. None of the 34 studies reviewed in Ward (2002) use data exclusively in the post 1992 period, and only a few used data that spanned into the early and mid 1990s. Second, instead of using a Bernoulli process to describe the dynamics of regime switching as in KGH (1993) and KG (1997), we employ the algorithm in Hamilton's regime switching model (1989) assuming a first order Markov process.[3] Finally, given the nature of fed cattle purchasing patterns, we use weekly data as opposed to KGH's use of daily data. We believe this provides a better (i.e., less noisy) platform for understanding the potential breakdowns in cooperative behavior over a long time series and should do a good job measuring accurately the length of differing regimes if such patterns are present.

The remainder of the paper is structured as follows. Section 2 presents the conceptual model of dynamic noncooperative pricing and game among the beef packers with complete but imperfect information. Section 3 discusses the econometric model as applied to the beef packing industry. Section 4 provides a description of the data and the estimation results. Conclusions and suggestions for future research are in Section 5.

[2]For example, Stiegert, Azzam and Brorsen (1993) found that unanticipated supply shocks led to breakdowns of average cost pricing strategies by beef packers. Xia and Sexton (2004) show that pricing contracts tied to spot markets add to the marginal cost of packers allowing them to bid lower for spot market cattle.

[3]Hamilton's method has been extensively used in a variety of macroeconomic analyses to explain regime patterns in business cycles, inflation, labor markets and others. Google Scholar finds over 3000 articles citing the the Hamilton (1989) paper. More recently, Markov regime switching methods has been employed in pricing models. For a recent example, see Chen and Forsyth's (2009) application on natural gas markets.

2 Conceptual Model

The conceptual framework follows from KGH (1993), which is based on the GP (1984) noncooperative game theoretic model using trigger strategies. Here, we begin by describing the model in the context of our analysis. We assume that n beef packers with complete but imperfect information bid for live cattle in repeated game framework. The assumptions of the model are:

1. Beef packers purchase an undifferentiated product — fed cattle from the regional cash market;
2. No exit or entry in the long run is considered in the game;
3. Beef packers understand the market structure well;
4. Beef packers cannot observe the pricing actions by others but obtain a perception about the degree price competition based on their own operating margin;
5. Beef packers are risk neutral and maximize their expected profit.

Profit of the i^{th} beef packer is given by:

$$\pi_i(p_{it}, p_{jt}, z_t) = (r_t - p_{it}k)y_{it}(p_{it}, p_{jt}, W_t, \xi_t) - c_i(z_t, y_{it}) \tag{1}$$

where p_{it} is the price for live cattle paid by the i^{th} packer at time t, p_{jt} is a vector of cattle prices paid by all other packers, r_t is the price of boxed beef, k is the inverse of the proportion of live animal converted to beef (cutout ratio), W_t is a vector of exogenous variables, y_{it} is the beef quantity the i^{th} packer produces from fed cattle and other inputs, ξ_t is a random term, c_i is the variable processing cost of the i^{th} beef packer and is a function of z_t, a vector of non-cattle variable input prices, and y_{it}. The set up of the variable processing cost in equation (1) is fundamentally different from the one in KGH (1993). In the KGH model, they used daily data and assumed that all costs other than the purchase of fed cattle were fixed. Our use of weekly data is suggestive of production process with flexibility in using other non-cattle inputs such as energy and labor. We assume that beef production is in fixed proportion to the live cattle slaughtered each week using the conversion ratio k.

Given the packer's own pricing strategy s_{it}, other packers' strategies s_{jt} and the discount rate δ, beef packer i maximizes the sum of current and discounted expected future profits:

$$V_i(s_t) = E\left[\sum_{t=0}^{\infty}(\frac{1}{1+\delta})^t \pi_i(s_{it}, s_{jt})\right] \tag{2}$$
$$i \neq j, i, j = 1, ..., n \text{ and } 0 < \delta < 1$$

In the case of a single decision game, firms choose the input purchase price of live cattle under noncooperation: one-shot Nash given by p''; or cooperation: collusion given by p'. So beef packers will cooperate when:

$$V_i(p') > V_i(p'') \text{ for all } i = 1, ..., n \tag{3}$$

In a repeated game, however, there is a profit incentive to secretly defect from the collusive outcome by raising price and securing greater input quantities. Following GP, a trigger strategy is used to deter such behavior. If a packer secretly increases the cattle price offer to $p^* > p^{'}$, all packers offer the single-period Nash price $p^{''} > p^*$. Therefore, under the threat of a trigger strategy, when collusive pricing is achieved, packers tend not to defect and market stabilizes. This occurs when the expected returns from cooperation are greater than the expected returns from defecting, which is formally given by:

$$V_i(p') > \pi_i(p^*) + \tfrac{1}{1+\delta} V_i(p'') \text{ for all firms} \tag{4}$$

The GP model is constructed on the premise that firm cannot directly observe each other's actions. However, they can observe their own margin level which, in our context, is the difference between the boxed-beef price and the fed-cattle price. Their pricing strategies each period would be dependent on their own observed margin in the previous periods. Therefore, when beef packers cannot observe each other's pricing behavior, they try to maximize their value function $V_i(s_t)$ subject to a trigger strategy:

$$S_{it} = \begin{cases} p' & \text{if } \mu < m_{t-1} \\ p'' & \text{if } \mu \geq m_{t-1} \text{ in the last } T-1 \text{ periods} \end{cases} \tag{5}$$

where μ is the trigger margin level, and m_{t-1} is the margin level in the previous period. If the beef packer's own observed margin in the previous period is greater than the trigger level μ, this packer offers a cooperative price p'. However, if the observed margin in the previous $T-1$ periods is less than μ, this packer offers a noncooperative price p''. In this way, the trigger strategy allows cooperation among beef packers on the equilibrium path because any firm that defects would expect to be punished by getting low profits for all $T-1$ future periods.

With the trigger strategy under consideration by all firms, the value function for the i^{th} packer starting in the cooperative phase is given by the sum of the current period collusive profit and the discounted expected future profits weighted by the occurrence probability of cooperation and noncooperation:

$$\begin{aligned} V_i(p') &= \pi_i(p') + Pr(\mu < m_t)\delta V_i(p') \\ &+ Pr(\mu \geq m_t)\left[\sum_{t=1}^{T-1}\left(\frac{1}{1+\delta}\right)^t \pi_i(p'') + \left(\frac{1}{1+\delta}\right)^T V_i(p')\right] \end{aligned} \tag{6}$$

Let $\Pr(\mu \geq m_t) = F$ where F is a distribution function. Rewriting equation (6) yields:

$$V_i(p') = \frac{(1+\delta)\pi_i(p'')}{\delta} + \frac{(1+\delta)^T(\pi_i(p') - \pi_i(p''))}{(1-\delta)^T - (1+\delta)^{T-1} + ((1+\delta)^{T-1} - 1)F} \tag{7}$$

Beef packers choose the price that maximizes the expected returns, so the interior solution to the first order condition of equation (7) is:

$$\begin{aligned} \frac{\partial V_i}{\partial s_i} &= \frac{\partial \pi_i(p')}{\partial s_i}(1+\delta)^T\left[(1+\delta)^T - (1+\delta)^{T-1} + ((1+\delta)^{T-1} - 1)F\right] \\ &+ [\pi_i(p') - \pi_i(p'')]\left((1+\delta)^{T-1} - 1\right)\frac{\partial F}{\partial s_i} f = 0 \end{aligned} \tag{8}$$

where f is the density function of F. The actions of beef packers are discontinuous: they aggressively purchase fed cattle while in the noncooperative state and offer a lower price for fed cattle while in the cooperative state. It is assumed that the detection of cheating behavior and the subsequent punishment occurs in a timely manner. For a cooperative equilibrium to exist in the multiple-period game, p', p'' and δ must satisfy the following condition:

$$V_i(p') > \pi_i(p^*) + \sum_{t=1}^{T-1} \left(\frac{1}{1+\delta}\right)^t \pi_i(p'') + \left(\frac{1}{1+\delta}\right)^T V_i(p') \tag{9}$$

Equation (9) implies that expected returns from cooperation are greater than the profits from cheating for one period plus $T-1$ periods of Nash profits. If we can find T, μ, p', and p'' to satisfy equations (8) and (9), a cooperative equilibrium can be sustained.

In our game, price wars are part of the equilibrium behavior because the fed-cattle supply is subject to random unobservable shocks and the packers' price offers are not observed by their competitors. When a low margin is observed, packers cannot tell if it is a consequence of a deviation from cooperative pricing by one of their rivals or if it is due to some other factor such as an unanticipated fed-cattle supply shock. Thus, the model portends a market condition in which possibly lengthy periods of high margins are observed followed by possibly lengthy periods in which low margins are observed. As the GP theory prescribes, the presence of trigger strategies provide implicit restraints for firms to bid conservatively for live cattle so as to "not upset the apple cart".

3 Econometric Model

In this section, we develop an empirical model that allows for discontinuous patterns (i.e., regime switching) between cooperative and noncooperative periods. As is discuss below, it is important to control for margin fluctuations that are not attributable firm conduct. Beef packers are assumed to choose input prices of live cattle to maximize profits. Thus, we begin with the first order condition of equation (1), which is:

$$\frac{\partial \pi_i}{\partial p_i} = (r - p_i k)[\frac{\partial y_i}{\partial p_i} + \sum_{j \neq i} \frac{\partial y_i}{\partial p_j}\frac{\partial p_j}{\partial p_i}] - k y_i - \frac{\partial c_i}{\partial y_i}[\frac{\partial y_i}{\partial p_i} + \sum_{j \neq i} \frac{\partial y_i}{\partial p_j}\frac{\partial p_j}{\partial p_i}] = 0 \tag{10}$$

Assume that the effect of the j^{th} firm's price on the i^{th} firm's fed-cattle purchase is smaller than the effect of its own price, and that firms are symmetric. Let $\frac{\partial y_i}{\partial p_i} = \delta$ ($\delta > 0$) and $\frac{\partial y_i}{\partial p_j} = -\frac{\delta}{q}$, where $q > 1$ is a constant. Equation (10) can be rewritten as:

$$(r - p_i k - mpc_i)[1 - \frac{\sum_{j \neq i} \frac{\partial p_j}{\partial p_i}}{q}]\delta = k y_i \tag{11}$$

where mpc_i is the firm i's marginal processing cost. Let $\frac{\sum_{j \neq i} \frac{\partial p_j}{\partial p_i}}{q} = \beta$ where β is the sum of reactions to firm i's purchase price p_i. When $\beta = 0$, firms offer the one-shot Nash price, which is consistent with noncooperative conduct. The cooperative conduct occurs with $0 < \beta < 1$. Summing equation (11) over n symmetric firms gives:

$$(r_t - p_t k - mpc_t)(1 - \beta)\delta = k y_t \tag{12}$$

Rewriting equation (12) yields the beef packer marketing margin:

$$m_t = r_t - p_t k = mpc_t + \frac{k y_t}{(1 - \beta)\delta} \tag{13}$$

Note that fluctuations in marginal processing costs are linked directly to movement in the marketing margin. Thus, it is important to control for such variations to limit its impact on the conduct term that assigns variation to specific regime. For econometric estimation, we assume a Generalized Leontief processing cost function given by:

$$C_t(y, w) = y_t(\phi_{11} w_{1t} + \phi_{22} w_{2t} + 2\phi_{12}\sqrt{w_{1t} w_{2t}}) + y_t^2(\phi_1 w_{1t} + \phi_2 w_{2t}) \tag{14}$$

where w_1 is the labor price and w_2 is the energy price. Marginal processing cost for beef packers is specified as:

$$mpc_t = \phi_{11} w_{1t} + \phi_{22} w_{2t} + 2\phi_{12}\sqrt{w_{1t} w_{2t}} + 2y_t(\phi_1 w_{1t} + \phi_2 w_{2t}) \tag{15}$$

Some discussion of the last term in equation (13) is necessary before proceeding. Unlike many industrial organization studies, our focus is not on evaluating certain benchmark conditions such as Bertrand or Stackleberg leader pricing. Our focus is on evaluating the market for regime switching between high levels and low levels of cooperative behavior. In equation (13), we let $k/(1-\beta) = \beta^s$ where β^s is the regime-dependent conduct parameter. Our model provides a unique estimate for β^s when the market is in a cooperative regime and when it is in a noncooperative regime. When $\beta > 0$ in either regime, it infers that market power is present. The model also suggests that if there are regime-switching patterns related to market power, the β^s in the cooperative phase will be larger than the β^s in the noncooperative regime.

Given that β^s captures regime-switching (if present), the fundamental question thus becomes what might trigger the market into different regimes. We are interested in understanding the role of supply shocks in this regard. Stiegert, Azzam, and Brorsen (1993) found that the unanticipated supply shocks were responsible in breaking down average cost pricing in beef packing using quarterly data from 1972-1986. However, the use of quarterly data does not allow a determination for whether or not such shocks led directly to changes in market regimes or whether or not any such regime changes were sustainable through time. In this paper, we focus on the potential for disruptions between the derived demand for processed beef and supply of live cattle as a driving force in regime switching activity. In particular, when slaughter is higher (lower) than what is anticipated, our hypothesis is that this may provide the impetus to switch from a cooperative (noncooperative) to a noncooperative (cooperative) regime. It is important to note that all supply disruptions may not lead to a switch in behavior. While we do not observe why actual slaughter differs from anticipate, our hypothesis is suggestive of derived demand driven results. Thus, our expectation is that when the downstream market demands more beef than can be easily delivered, packers will become aggressive in attracting additional supply to their plants. In doing so, they would signal to other beef packers a breakdown in the cooperative regime. To integrate a design of this nature into our model, let δ reference the ratio of actual slaughter to predicted slaughter, i.e., $\delta = y_t/\hat{y}_t$. Using the marginal processing cost specification and the specification for regime switching, the marketing margin in equation (13) is now written as:

$$m_t = r_t - p_t k = \phi_{11} w_{1t} + \phi_{22} w_{2t} + 2\phi_{12}\sqrt{w_{1t} w_{2t}} + 2y_t(\phi_1 w_{1t} + \phi_2 w_{2t}) + \beta^s \hat{y}_t \qquad (16)$$

We obtain anticipated supply in a manner similar to that used by Stiegert et al. (1993) but modified for weekly data. Specifically, we estimated the following auxiliary regression model:

$$y_t = \alpha + \alpha_1 y_{t-1} + \alpha_2 \mathrm{cof}_1 + \alpha_3 p_c + \sum_{i=4}^{6} \alpha_i \mathrm{plc}_i + \sum_{i=7}^{17} \alpha_i D_i + e_t \tag{17}$$

where y_{t-1} is slaughter in the previous week, cof_1 is the previous month's cattle on feed, p_c is the price of corn, plc_i are cattle placements lagged 4, 5, and 6 months, and the final term contains 11 monthly dummy variables. We assume that predicted slaughter is estimated by $\hat{y}_t$ for each week: $y_t - e_t = \hat{y}_t$. The auxiliary regression is estimated using GLS.

Weekly changes in the marketing margin is a recurrent feature of cattle markets. Our empirical model will look to evaluate that part of the marketing margin that is not related to marginal processing costs for potential regime switching behavior. While the conduct associated with either cooperation or noncooperation is unobserved, the potential for switching between each regime is modeled using a Multivariate Markov-Switching framework given by:

$$m_t = \nu_{st} + \beta^s \hat{y}_t + \gamma_1 w_{1t} + \gamma_2 w_{2t} + \gamma_3(2\sqrt{w_{1t}w_{2t}}) + \gamma_4(2y_t w_{1t}) + \gamma_5(2y_t w_{2t}) + \varepsilon_t \tag{18a}$$

$$\nu_{st} = \kappa_1 \xi_t^1 + \kappa_2 \xi_t^2 \tag{18b}$$

$$\varepsilon_t | S_t \sim N(0, \sigma_{st}^2) \text{ where } \sigma_{st}^2 = \rho_1 \xi_t^1 + \rho_2 \xi_t^2 \tag{18c}$$

Let $S_t = \{1, 2\}$ denote the 2-state unobserved regimes with $S_t = 1$, representing the noncooperative regime and $S_t = 2$, representing the cooperative regime. The transition between these two states is governed by a first-order Markov process:

$$Prob[S_t = 1 | S_{t-1} = 1] = p \qquad Prob[S_t = 2 | S_{t-1} = 1] = 1 - p \tag{19a}$$

$$Prob[S_t = 2 | S_{t-1} = 2] = q \qquad Prob[S_t = 1 | S_{t-1} = 2] = 1 - q \tag{19b}$$

Following Bellone (2005), define the time dependent information sets I_t available in each regime, which are equal to the shadow random variables in equations (18a) and (18b): $\xi_t^1 = I_{S_t=1}$ and $\xi_t^2 = I_{S_t=2}$. As a result, we can define the conditional probabilities related to each state as:

$$P(S_t = 1|I_t) = E(\xi_t^1|I_t) \tag{20a}$$

$$P(S_t = 2|I_t) = E(\xi_t^2|I_t) \tag{20b}$$

In the Markov-Switching model defined in (18a), besides the intercept ν_{st}, $\hat{y}_t$ is the only exogenous variable that is subject to switching regimes. All other terms are related to marginal processing costs determined after the non-cattle input price and subsequent supply levels are obtained in the market. The error terms are assumed to have different variances in the two regimes. Therefore, $(\beta^s, \kappa_1, \kappa_2, \rho_1, \rho_2)$ is the vector of regression coefficients which are regime-dependent, and $\gamma = (\gamma_1, \ldots, \gamma_5)$ is the vector of regression coefficients which are regime-independent.

Following the estimation of the Multivariate Markov-Switching models developed by Bellone (2004, 2005), with the normality assumption of ε_t, the conditional probability density function of m_t is given by:

$$f(m_t|S_t = j, I_{t-1}, \Theta) = \frac{|\Sigma_j^{-1/2}|}{(2\pi)^3} exp(-\frac{\varepsilon_t' \Sigma_j^{-1} \varepsilon_t}{2}) \tag{21}$$

where $\Theta = (p, q, \beta^s, \gamma, \kappa, \rho)$ and $\Sigma_j = \rho_1 \xi_t^1 + \rho_2 \xi_t^2$. Then the unconditional density of m_t is calculated by summing conditional densities over the two values of S_t:

$$f(m_t|I_{t-1}, \Theta) = \Sigma_{j=1}^2 P(S_t = j|I_{t-1}, \Theta) f(m_t|S_t = j, I_{t-1}, \Theta) \tag{22}$$

The maximum likelihood estimate of Θ is obtained by maximizing the following log likelihood function:

$$L(\Theta) = \Sigma_{t=1}^T ln(f(m_t|I_{t-1}, \Theta)) \tag{23}$$

4 Estimation Procedures and Data

The data sets used in this paper were collected from Livestock Marketing Information Center (LMIC), National Agricultural Statistics Service (NASS) and the Department of Labor. One analysis was conducted for the national beef market and one for state of Kansas. Kansas represents a prominent beef production region where cattle feeding and slaughter industries are both present. The weekly slaughter, cattle placement, cattle on feed, and corn price data were used in the auxiliary regression described in equation (17). The Markov switching regression used processing margins and beef processing cost data. The processing margins were obtained by subtracting from the boxed beef price the regional or national fed-cattle price converted to a carcass equivalent (price/0.615). To remove the impact of inflation, the margin values are deflated to a 1992 base year. The energy price index is from the producer price index for the meat packing industry and the labor price is the average hourly production worker earnings for the meat packing industry.

The study period for this analyzed was selected for several reasons. First, stability in industry structure is generally preferred in the estimation of the GP models. This allows for a more confident interpretation of regime switching behavior because it limits the possibility of assigning switching patterns to relatively more permanent structural change. The period from February 1992 to January 2000 represents a good time frame for analysis. In the years previous to 1992, the market had undergone significant restructuring through mergers and plant closures. The HHI passed the DOJ threshold of 1800 in 1992 but then leveled off in future years. After 2000, the beef market may have been significantly impacted by additional restructuring via mergers, increased regulations related to food safety concerns and bioterrorism concerns after terrorists attacks on 9/11/2001.

The results from the estimation of equation (17) are reported Table I. Not surprising, the coefficient on slaughter lagged one week was significant for both Kansas and the national market. It appears to capture trends in slaughter that relate to weather or other factors in the normal animal production process. Cattle placements were not statistically significant in either regression. It is plausible that lagged slaughter, lagged cattle on feed and seasonal variables explain much of the variation in slaughter that placement data might otherwise capture. Additionally, multicollinearity within the placement variables and the corn price could be present. A joint test of the three lagged placements and corn price was highly statistically significant: F-stat for Kansas=3.77 and F-stat for the national market=3.20. Both are larger than the 1% critical value of 3.17. This indicates that there is important information in this group of variables useful for predicting slaughter on a weekly basis. The adjusted R^2 numbers for Kansas and the U.S. indicate that the regressions do a good job of predicting slaughter.

We next estimated the Markov-Switching model described by equation (18) using MSVAR-lib developed by Bellone (2005). We let state 1 be the noncooperative state and state 2 be the cooperative state. Applying Hamilton's (1989) algorithms of filtering and smoothing to margin changes in one regional market and the national market, numerical maximization of the conditional log likelihood function led to the maximum likelihood estimates of the transition probabilities, and the regime dependent and independent parameters.

The regression estimates are presented in Table II. The first section of Table II contains the estimated regime-independent parameter values associated with marginal processing costs. Almost all the γ terms are statistically significant in both regressions, which means the marginal processing cost components are important in explaining weekly movements in the margin.

The next section of Table II contains the estimates for each regime-dependent parameter. The conditional probabilities of remaining in the noncooperative regime (p) and the cooperative regime (q) are estimated through the EM algorithm and reported first. Low (high) conditional probabilities suggest that switching between regimes is both easy (difficult) and frequent (infrequent). Importantly, all of these estimates are highly statistically significant and in a range between 0.936 and 0.970. This indicates that the markets analyzed have a high probability of remaining in the previous period regime. Using the Kansas results as an example, the term prob$(S_t = 2|S_{t-1} = 2) = q = 0.936$ means that there is a 93.6% chance that packers will cooperate when they cooperate in the previous week. Similarly, there is a 96.9% chance that Kansas packers will remain in a noncooperative state when they are in a noncooperative state the previous week. These findings suggest that it not be easy for the market to switch regimes.

Both of the β^s parameters for each regression are positive and are all statistically significant. Additionally, each of the β^s parameters in the cooperative regime is quite a bit larger in magnitude than the comparative β^s in the noncooperative regime. This infers that the firms observe the different regimes and switch their behavior to comply to either a more cooperative or a more competitive environment. The intercept is a regime-dependent parameter controlled by κ_1 and κ_2. All are significant in both regressions. Finally, the variances of each regression are allowed to vary depending on the regime and controlled through ρ_1 and ρ_2. The estimates of ρ_1 and ρ_2 provide clear evidence that the variances differ strongly in each regime in both regressions. The estimates and the significance of the $\rho's$ and the $\kappa's$ offer additional support for the presence of a switching pattern.

The focus of the remaining analysis was on defining and analyzing the conditional outcomes associated with switching. Four analyses were conducted. First, as shown at the bottom of Table II, we calculated the expected duration of cooperation and noncooperation respectively. Conditional on being either in cooperative state or noncooperative state, the expected durations are calculated by:

$$\sum_{\lambda=1}^{\infty} \lambda p^{\lambda-1}(1-p) = (1-p)^{-1} \tag{24a}$$

$$\sum_{\lambda=1}^{\infty} \lambda q^{\lambda-1}(1-q) = (1-q)^{-1} \tag{24b}$$

The expected duration of cooperation is about 15.63 weeks for Kansas and 21.28 weeks for the national market. The expected duration of noncooperation is about 33 weeks for Kansas and the national market. Note that Kansas appears representative of the national market during the noncooperative regime. The duration of the cooperative phase in Kansas is about 25% lower compared to the national market. This is likely the regional nature of the Kansas regression. Although Kansas is central to the largest feeding and slaughtering region in the U.S., supply injections and leakages to and from neighboring states could disrupt the status quo and thus cause localized breakdowns in cooperative regimes. Clearly, spatial arbitrage opportunities would require that a region eventually align its pricing with other regions. We develop this subject further in the second and third analyses of switching patterns.

Second, using Hamilton's (1989) filter techniques, we calculate the probability of being in a state of cooperation ($S_t = 2$) or noncooperation ($S_t = 1$) at time t. We present the week-to-week probabilities of cooperation for each market in Figure 1. Note that the unshaded regions represent the weeks in noncooperative states and the shaded regions represent weeks in cooperative states.[4] As we can see in the figure, there are 8 cooperative periods for Kansas and 6 for the national market during the study years. Note also that the cooperative regime for the U.S. centered on week 225 is broken into two very short cooperative regimes in Kansas. It appears the Hamilton algorithm could not keep Kansas in a single regime during these weeks. This helps to explain why the average cooperative duration for Kansas is 25% shorter than for the national market.

Third, we investigated the conditions that lead to each of the 16 switches in Kansas and 12 switches for the U.S. market. Our model setup suggests that shocks in the actual versus anticipated slaughter levels are a potential source of the switching behavior. In Table III, we report the percentage of unanticipated slaughter during the week that a switch takes place. Note that for the U.S. market, there are six weeks with a switch to a noncooperative state (unshaded rows) and six weeks with a switch related to a cooperative state (shaded rows). The results in Table III reveal an interesting pattern of unanticipated slaughter related switching behavior. Specifically, we note that a breakdown of cooperative behavior is strongly consistent with slaughter levels above the anticipated level. For Kansas, in each of the weeks that the market switched to a noncooperative state, unanticipated slaughter is 1.51%-10.26% above the anticipated level. For the national market, unanticipated supply is 1.12%-11.68% above the level of anticipated slaughter in the weeks that cooperation breaks down. The results suggest that when packers need larger levels of supply than are, perhaps, readily available, cooperation breaks down. The reverse pattern is similar when the market switches to a cooperative regime. The actual slaughter is below anticipated slaughter in all switches in the Kansas market (-0.44% to -8.08%) and for the national market (-1.17% to -4.66%). These results suggest that when slaughter is below the anticipated level, beef packers recognize that supplies have become plentiful and switch to less aggressive bidding stance. This softer bidding has the potential to last many weeks past the event.

Fourth, using the shaded and unshaded regions presented in Figure I, we calculated the average regime dependent marketing margins. These results are in Table IV. Our results for the national market indicate that the cooperative regime produces a marketing margin

[4] Following Hamilton (1989), our decision rule is that beef packers are in the cooperative regime when $\mathrm{P}[S_t = 2] > 0.5$, and they are in the noncooperative regime when $\mathrm{P}[S_t = 1] > 0.5$.

that is 65.76% above the marketing margin for the noncooperative regime. Kansas is slightly lower at 60.26%. Based on expected regime lengths of 21 weeks cooperative and 33 weeks noncooperative; on average, the national market is in a cooperative state 39% of the time. This means that the expected increase in the national marketing margin due to switching patterns is 25.65%. For Kansas the result is a 19.64% increase in the margin. Based on annual production of beef in the national market over the study years, a 25.65% increase in the marketing margin infers an increase in profits to the packing industry of nearly 491 million 1992 dollars each year.[5] Additionally, this shift in rent does not account for profits that might occur due to market power in the noncooperative regime. The β^s estimates are significant in the noncooperative regime which indicates the presence of a lower level of market power compared to the cooperative period.

5 Conclusion

In 1992, concentration in the U.S. beef packing industry for the first time passed over the HHI=1800 level, which is recognized by the Department of Justice as the threshold for a heavily concentrated industry. Most previous studies of market power in this industry analyzed data in periods prior to 1992. In this article, we construct a GP model using Hamilton's Markov-Regime-Switching technique to analyze weekly marketing margin data for the U.S. beef packing industry from 1992-2000. We focus on changes in the normal throughput of supply as a potential catalyst for regime switching behavior. The model was estimated for one region (Kansas) and for the national market. Both models produced results with sufficient statistical support for switching behavior. Each estimation identified similar cooperative periods and noncooperative periods that mostly began and ended within 1-4 weeks of each other.

A major finding from the study was that slaughter levels higher than anticipated led to breakdowns in cooperative regimes while lower than anticipated slaughter provided a way for packers to switch to a cooperative regime. This seems to suggest that cooperation,

[5] From 1992 to 2000, annual beef production in the U.S. was in a range primarily from 25-27 billion lbs. For 26 billion pounds of production: $(26,000,000,000/100) * (7.36 - 2.52) * 0.39 =$490.78mil.

once established as a norm, can last until insufficient short-term supplies spark a return to more competitive behavior. Our analysis indicates that cooperative regimes are expected to last about 21 weeks while noncooperative regimes last about 33 weeks. During the study years, national concentration levels remained fairly stable, so regimes patterns were likely not unduly influenced by major structural changes to the industry. Our results suggest that beef packers have market power that is manifested in ways that would not be easily detected in more aggregated data or with a model that does not provide sufficient flexibility in allowing for major breakdowns and returns to cooperation. Since 2000, the concentration levels in beef packing have continued to rise. It is likely that such changes provide beef packers with even more ability to lengthen cooperative phases and extract greater rents than during our study years.

References

[1] Azzam, A. and D. Anderson (1996) Assessing Competition in Meatpacking: Economic History, Theory, and Evidence, *GIPSA-RR 966*, May.

[2] Azzam, Azzeddine M. and T. Park (1993) Testing for Switching Market Conduct, *Applied Economics; 25*, 795-800.

[3] Barkema, Alan, M. Drabenstott and N. Novack (2001) The New U.S. Meat Industry, *Federal Reserve Bank of Kansas City/ Economic Review; Second Quarter 2001*, http://www.kansascityfed.org/PUBLICAT/ECONREV/PDF/2q01bark.pdf

[4] Bellone, Benoit (2004) MSVARlib: A New Gauss Library to Estimate Multivariate Hidden Markov Models.

[5] Bellone, Benoit (2005) Classical Estimation of Multivariate Markov-Switching Models using MSVARlib.

[6] Chen, Zhuliang, and Peter A. Forsyth (2009) Implications of a Regime-Switching Model on Natural Gas Storage Valuation and Optimal Operation, *Quantitative Finance, January 2009*, 1-18.

[7] Grain Inspection, Packers and Stockyards Administration (2002) Packers and Stockyards Statisitical Report 2002, *U.S. Department of Agriculture, Washington, D.C.*

[8] Green, E.J. and R.H. Porter (1984) Noncooperative Collusion Under Imperfect Price Information, *Econometrica; 52(1)*, 87-100.

[9] Hamilton, James D. (1989) A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle, *Econometrica; 57(2)*, 357-84.

[10] Koontz, Stephen R., Philip Garcia and Michael A. Hudson (1993) Meatpacker Conduct in Fed Cattle Pricing: An Investigation of Oligopsony Power, *American Journal of Agricultural Economics; 75*, 537-48.

[11] Koontz, Stephen R. and Philip Garcia (1997) Meat-Packer Conduct in Fed Cattle Pricing: Multiple-Market Oligopsony Power, *Journal of Agricultural and Resource Economics; 22(1)*, 87-103.

[12] Sexton, R.J. (2000) Industrialization and Consolidation in the U.S. Food Sector: Implications for Competition and Welfare, *American Journal of Agricultural Economics; 82:5*, 1087-1104.

[13] Ward, Clement E. and Ted C. Schroeder (2002) Structural Changes in Cattle Feeding and Meat Packing, *Oklahoma Cooperative Extension Service AGEC-553.* http://pods.dasnr.okstate.edu/docushare/dsweb/Get/Document-2009/AGEC-553web.pdf

[14] Ward, Clement E. (2002) A Review of Causes for and Consequences of Economic Concentration in the U.S. Meatpacking Industry, *Current Agriculture, Food and Resource Issues; 3*, 1-28.

[15] Xia, Tian and Richard Sexton (2004) A Review of Causes for and Consequences of Economic Concentration in the U.S. Meatpacking Industry, *American Journal of Agricultural Economics, 86(1)*, 124-138.

[16] Stiegert, Kyle W., Azzeddine Azzam, and B.Wade Brorsen (1993) Markdown Pricing and Cattle Supply in the Beef Packing Industry, *American Journal of Agricultural Economics; 75*, 549-58.

[17] Whinston, Michael D. (2007) Antitrust Policy toward Horizontal Mergers, *Handbook of Industrial Organization; 3.*

Table I: GLS Estimates of Weekly Marketing Model

	Kansas		U.S.	
	estimate	s.d.	estimate	s.d.
cons.	9.479***	(1.206)	17.803***	(2.282)
y_{t-1}	0.748***	(0.027)	0.726***	(0.028)
cof_{-1}	0.073*	(0.039)	0.066***	(0.017)
p_c	-0.120	(0.111)	0.173	(0.212)
plc_{-4}	0.133	(0.134)	0.040	(0.072)
plc_{-5}	0.113	(0.150)	-0.012	(0.089)
plc_{-6}	0.073	(0.120)	0.045	(0.065)
D_2	-1.026***	(0.345)	-2.205***	(0.611)
D_3	-1.185***	(0.347)	-2.214***	(0.797)
D_4	0.032	(0.380)	-0.318	(0.914)
D_5	1.480***	(0.380)	2.762***	(0.733)
D_6	0.984**	(0.404)	2.287***	(0.762)
D_7	0.153	(0.377)	0.469	(0.712)
D_8	1.037***	(0.367)	2.022***	(0.714)
D_9	0.126	(0.390)	0.676	(0.750)
D_{10}	-0.172	(0.364)	-0.183	(0.740)
D_{11}	-0.366	(0.388)	-0.730	(0.739)
D_{12}	-1.922***	(0.374)	-3.495***	(0.608)
Adjusted R^2	0.823		0.833	

*** $p < 0.01$, **$p < 0.05$, * $p < 0.10$.

Table II: Maximum Likelihood Estimates

	Parameters	Kansas		U.S.	
		estimate	s.d.	estimate	s.d.
Regime-Independent MLE	γ_1	-4.244***	(0.952)	-3.505***	(1.005)
	γ_2	-1.800***	(0.435)	-0.887*	(0.465)
	γ_3	3.594***	(0.793)	2.609***	(0.839)
	γ_4	1.006*	(0.549)	1.581***	(0.550)
	γ_5	-0.270	(0.312)	-0.774***	(0.318)
Regime Dependent MLE	p	0.969***	(0.012)	0.970***	(0.014)
	q	0.936***	(0.024)	0.953***	(0.015)
	β^1	0.131***	(0.056)	0.082*	(0.045)
	β^2	0.239***	(0.078)	0.268***	(0.074)
	κ_1	-0.417***	(0.034)	-0.561***	(0.035)
	κ_2	0.941***	(0.069)	0.820***	(0.046)
	ρ_1	0.203***	(0.020)	0.204***	(0.015)
	ρ_2	0.317***	(0.046)	0.329***	(0.034)
Expected Duration	Non-Cooperation	32.26		33.33	
	Cooperation	15.63		21.28	

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table III: Percentage of Unanticipated Slaughter in the Switching Weeks

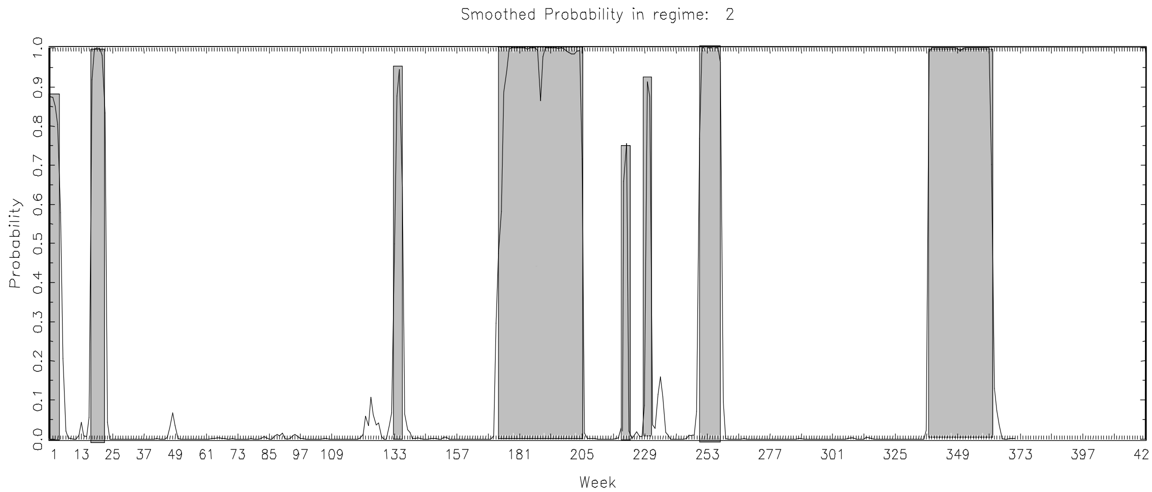
Kansas		U.S.	
Week	%	Week	%
2/1/1992	-2.58		
3/21/1992	3.15		
5/9/1992	-0.44	4/11/1992	-2.01
7/18 /1992	2.33	5/16/1992	4.52
7/30/1994	-2.93	4/16/1994	-1.54
9/24/1994	2.43	9/3/1994	3.23
4/22/1995	-2.09	4/8/1995	-5.55
12/16/1995	1.51	12/9/1995	5.76
4/20/1996	-2.36	3/30/1996	-4.66
5/11/1996	3.02		
6/22/1996	-8.08		
7/6/1996	2.44	6/1/1996	5.10
11/16/1996	-3.47	10/26/1996	-1.17
1/4/1997	10.26	12/14/1996	1.12
7/18/1998	-2.62	6/6/1998	-3.38
1/16/1999	4.70	1/2/1999	11.68

Rows highlighted in gray represent switches into the cooperative regime.
Rows not highlighted represent switches into non-cooperative regimes.
Percentage change results are from $(y-\hat{y})/\hat{y}$.

Table IV: Margins in Cooperative and Noncooperative Regimes

Region	Cooperative Margin	Noncooperative Margin	% Increase in the Cooperative Regime
Kansas	6.92	2.75	60.26
U.S.	7.36	2.52	65.76

Unit: $/hundredweight boxed beef



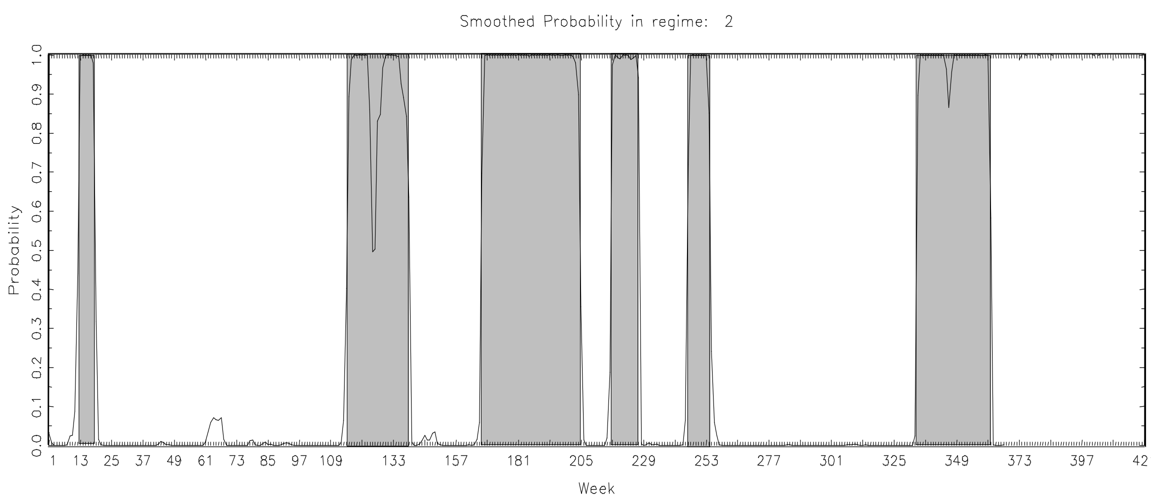


Figure 1: Probability of cooperative regime for Kansas (above graph), and U.S. (lower graph). Shaded regions define periods of cooperative regime.