Implications of Increased Regional Concentration and Oligopsonistic Coordination in the Beef Packing Industry

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This article proposes an oligopsony pricing model for projecting the effects of increased concentration or oligopsonistic coordination in beef packing using simulation methods. The model combines an explicit behavioral theory of packing firms with an attempt to respect the regional scope of cattle procurement markets. Our results indicate less danger of falling cattle prices, as a result of increased packer concentration or coordination, than do results from conventional econometric studies.

Key words: beef packing, conjectural variation, oligopsony, regional concentration, simulation.

Concern about the impact on live cattle prices of increased concentration in the beef packing industry has prompted several empirical studies. Notable examples of relevant econometric work in the Bain tradition [hereafter, the "structure-conduct-performance (or SCP) approach"] include Ward (1981, 1985); Menkhaus, St. Clair, and Ahmaddaud; and Quail et al.[1] Other work, including Azzam and Pagoulatos; Schroeter (1988); and Schroeter and Azzam, uses the conjectural variation approach to modeling equilibrium in homogeneous product oligopoly and tests for packer price-taking behavior in both factor and output markets.[2] Both approaches have the same objective: to

determine how recent (or future) increases in concentration have affected (or will affect) cattle prices. Each approach has its shortcomings, however.

The SCP approach seeks to infer the degree of competition in cattle procurement markets through ad hoc models relating a performance measure (live cattle price) to structural characteristics (including concentration) using regional data. The structural characteristics in the empirical models are supplemented by other components believed by the investigator to affect performance. The problem is that the models are not explicitly connected to behavior at the firm level. To begin, the manner in which variables are chosen for use in estimation is often "more akin to a literature search for a list of possible variables than the development of a coherent structure-performance model" (Sawyer, p. 296). More fundamentally, the models and estimation procedures fail to test or impose important restrictions implied by theory (Schroeter 1990). As a result, the empirical estimates from the SCP approach come with no guarantee of consistency with a coherent model of rational firm conduct.

The conjectural variation approach, on the other hand, is explicitly theory based. "The behavioral equations by which firms set price

and quantity (are) . . . estimated, and parameters of those equations . . . directly linked to analytical notions of firm and industry conduct" (Bresnahan, p. 1012). The problem here is that, to date, applications of the conjectural variation approach to the problem of testing competitive behavior in the beef industry have not adequately addressed the fact that relevant cattle procurement markets are regional, not national, in scope.

This article proposes an alternative, non-econometric, procedure for projecting the price effects of increased concentration or oligopsonistic coordination in beef packing. It combines an explicit behavioral theory of beef packing firms with an attempt to respect the regional scope of competition in cattle procurement. The behavioral model is used to derive the relationships among structure, firm conduct, and performance in cattle markets. Not all of the current values of the structure/conduct/performance parameters are directly observable, however, so the model first must be calibrated using estimates of these before it can serve as a basis for simulations of the performance effects of future changes in structure or conduct. Our simulation results suggest somewhat less danger of falling cattle prices due to increasing concentration than do the results of conventional econometric studies utilizing the SCP approach.

A Model of Oligopsony Pricing in Regional Cattle Markets

Let K denote the number of cattle procurement regions. Each packer in every region is assumed to be a perfect competitor in a common national output market with price p.[3] Denote the annual output of packer j in region k by q_{kj}. Adopting the convention of measuring cattle inputs in carcass weight equivalent units, q_{kj} also represents the annual livestock input of packer j in region k. The market inverse cattle supply function in region k is given by $w_k = g(Q_k)$, where w_k is the price of cattle in region k, $Q_k = \sum_{j=1}^{N_k} q_{kj}$ is total cattle sales volume

for the year in region k, and N_k is the number of packers in region k. If the jth packer in the kth region has processing cost function $C_{kj}(q_{kj})$, then the packer's annual profit is given by

$$\pi_{kj} = (p - w_k)q_{kj} - C_{kj}(q_{kj}).$$

Packers choose their cattle input quantities to maximize profit. A firm with market power will internalize the effect that its choice of quantity will have on regional quantity and, in turn, on regional cattle price. The first-order condition for profit maximization is obtained by differentiating the profit function with respect to q_{kj} and setting the result equal to zero, yielding

(1) $p - C'_{kj}(q_{kj}) - w_k = q_{kj}\left(\dfrac{dw_k}{dQ_k}\right)(1 + \lambda_{kj}).$

Here, $\lambda_{kj} = \dfrac{d \sum\limits_{i \neq j} q_{ki}}{dq_{kj}}$ and can be interpreted as the jth firm's conjecture regarding its regional rivals' responses to a change in its own input. In effect, λ_{kj} identified the jth firm's degree of market power. But rather than work with the firm-specific λ_{kj}, we follow Clarke and Davies in adopting a parametrization of them that leads to region-specific conduct indices. Assume that the behavior of each firm in region k is consistent with the expectation that its rivals' proportionate quantity responses will all be a constant multiple, say α_k, of its own proportionate change. That is,

$$\frac{dq_{ki}}{q_{ki}} = \alpha_k\left(\frac{dq_{kj}}{q_{kj}}\right) \qquad \text{for all } i \neq j$$

or

$$\frac{dq_{ki}}{dq_{kj}} = \alpha_k\left(\frac{q_{ki}}{q_{kj}}\right) \qquad \text{for all } i \neq j.$$

Substituting this expression into the definition of λ_{kj} yields

(2) $\lambda_{kj} = \alpha_k\left(\dfrac{Q_k}{q_{kj}} - 1\right).$

Defined this way, α_k provides an index, with values between zero and one, of the degree of firms' implicit coordination in the kth regional cattle market. A value of zero for α_k can be identified with firms anticipating no responses to their output adjustments. This is Cournot conduct, the polar case of noncooperative behavior. When α_k is set to one, equations (2) and (1) imply that output price minus marginal

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[3] This may not be a tenable assumption. However, since most of the attention seems to have focused on market power in procurement markets rather than output markets, we proceed accordingly.

processing cost is equal to marginal factor cost computed from the region's aggregate cattle supply curve. Thus, $\alpha_k = 1$ signals perfect cattle market collusion among region k's packers.[4]

Substituting (2) into (1) and multiplying by the jth firm's market share yields

$$(p - w_k)\left(\frac{q_{kj}}{Q_k}\right) - \frac{q_{kj}C'_{kj}(q_{kj})}{Q_k}$$
$$= Q_k\left(\frac{dw_k}{dQ_k}\right)\left\{\left(\frac{q_{kj}}{Q_k}\right)^2 + \alpha_k\left[\left(\frac{q_{kj}}{Q_k}\right) - \left(\frac{q_{kj}}{Q_k}\right)^2\right]\right\}.$$

Summing over firms in the kth region and dividing by w_k yields

$$\frac{p - c_k - w_k}{w_k} = \frac{[\alpha_k - H_k(\alpha_k - 1)]}{\epsilon_k},$$

where c_k is a market-share weighted average of region k's packers' marginal processing costs, ϵ_k is the elasticity of cattle supply in region k, and $H_k = \sum_{j=1}^{N_k}\left(\frac{q_{kj}}{Q_k}\right)^2$ is the Herfindahl index of concentration in region k.

Under the assumption of a competitive output market, $p - c_k$ is the value of the marginal product of cattle net of packers' average marginal processing costs. Since this would equal the factor price under competition, $(p - c_k - w_k)/w_k$ provides a relative measure of the oligopsony distortion in region k's cattle market. Denoting this by D_k, we have

$$D_k = \frac{[\alpha_k - H_k(\alpha_k - 1)]}{\epsilon_k}.$$

Notice that, for given values of the conduct parameter, α_k, and supply elasticity, ϵ_k, the distortion increases with concentration, H_k. Thus increasing concentration worsens performance even with completely noncooperative, or Cournot, conduct. For given conduct and concentration, the distortion decreases as supply elasticity increases. Thus even tightly coordinated oligopsony becomes less distortionary as

supply responsiveness grows. Finally, for a given concentration and supply elasticity, the distortion increases as α_k increases and approaches the pure monopsony level, $(1/\epsilon_k)$, as α_k approaches the value of one consistent with perfect collusion. Specifically, when $\alpha_k = 1$, H_k drops out of the expression: In a joint profit-maximizing input buyers' cartel, market performance is independent of buyer concentration. From (3), a quantity weighted average of regional oligopsony price distortions can be formed as

$$(4) \qquad D = \sum_{k=1}^{K} D_k \frac{Q_k}{Q}$$
$$= \sum_{k=1}^{K}\left(\frac{\alpha_k - H_k(\alpha_k - 1)}{\epsilon_k}\right)\frac{Q_k}{Q},$$

where $Q = \sum_{k=1}^{K} Q_k$ is national cattle sales volume.

The starting point of each simulation reported in the next section is a "baseline case" which combines estimates of regional quantities and concentration indices, $\hat{Q}_k$s and $\hat{H}_k$s, with assumed values for the current conduct parameters, $\hat{\alpha}_k$s, and an estimate of the national average oligopsony distortion, $\hat{D}$, formed from industry financial data. Each baseline case is a candidate description of the current situation in the industry. Assuming that supply elasticities are constant across regions, equation (4) can be used to solve for the common value of ϵ that is consistent with any particular baseline case:

$$(5) \qquad \hat{\epsilon} = \frac{1}{\hat{D}\hat{Q}} \sum_{k=1}^{K}\left[\hat{\alpha}_k - \hat{H}_k(\hat{\alpha}_k - 1)\right]\hat{Q}_k,$$

where $\hat{Q} = \sum_{j=1}^{K} \hat{Q}_k.$

Establishing a baseline case enables comparison with "test cases." These will be characterized by sets of prospective future values for concentration and conduct parameters, H_k^*s, and α_k^*s. Assuming that regional supply elasticities will be unaffected by changes in packer concentration or conduct, the regional and national average price distortions corresponding to the test case are

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[4] The collusion we have in mind is implicit collusion. As Posner (p. 40) explained, "in some circumstances competing [firms] might be able to coordinate their pricing without conspiring in the usual sense of the term—that is, without any overt or detectable acts of communication. This is the phenomenon that lawyers call conscious parallelism and some economists call oligopolistic interdependence, but I prefer to call implicit collusion in contrast to explicit collusion of the formal cartel or its underground counterpart."

(6)
$$D_k^* = \frac{\alpha_k^* - H_k^*(\alpha_k^* - 1)}{\hat{\epsilon}}$$

and

(7)
$$D^* = \sum_{k=1}^{K} D_k^* \frac{\hat{Q}_k}{\hat{Q}}.$$

D^* is the estimate of the national average oligopsony distortion that would result due to the test case's hypothetical changes in concentration and/or conduct.[5] The corresponding effects on national cattle sales volume and average price can be projected given assumptions about the aggregate (that is, national) supply and demand curves.[6] Assume constant elasticity forms for the national market's cattle supply curve,

(8)
$$Q = bw^e, \quad e > 0,$$

and derived demand curve

(9)
$$Q = dw^\eta, \quad \eta < 0.$$

At this point, normalize by setting current, or baseline case, values of price and quantity to 1 and 100, respectively.[7] As figure 1 illustrates, the supply price at $\hat{Q} = 100$ will then be $\hat{w} = 1$ while the derived demand price at this quantity will be $(1 + \hat{D})$, where $\hat{D}$ is the estimate of the current distortion. Substituting these values into equations (8) and (9) and solving for b and d yields

$$b = 100 \quad \text{and} \quad d = 100(1 + \hat{D})^{-\eta}.$$

Denote the test case projections of the national market quantity and average price by Q^* and w^*, respectively. Again referring to figure 1, the point with these coordinates must lie on the supply curve:

$$Q^* = 100(w^*)^e.$$

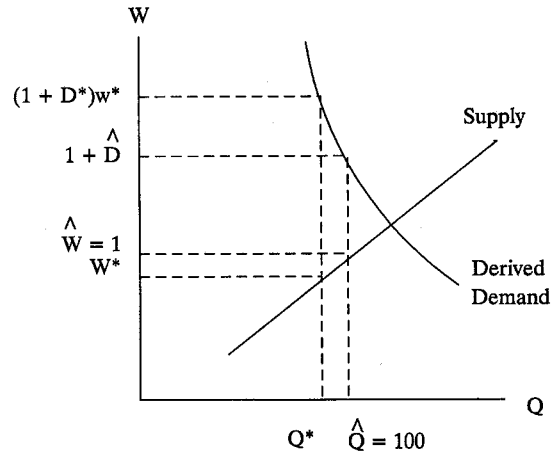Moreover, since Q^* and w^* must be consistent with an oligopsony distortion of D^*, the de-

Figure 1. A model for calculating price and quantity impacts of changing structure and conduct

mand price at Q^* must be $w^*(1 + D^*)$:

$$Q^* = 100(1 + \hat{D})^{-\eta}[w^*(1 + D^*)]^\eta.$$

Solving the last two equations for Q^* and w^* yields

(10)
$$Q^* = 100\left[\frac{1 + D^*}{1 + \hat{D}}\right]^{\frac{\eta e}{e - \eta}}$$

and

(11)
$$w^* = \left[\frac{1 + D^*}{1 + \hat{D}}\right]^{\frac{\eta}{e - \eta}}.$$

Simulations

The first step in performing the simulations is to establish the parameters of the baseline case. Bruce Marion provided 1986 data for the regional Herfindahl indices and annual slaughter volumes for the 13 fed cattle marketing regions described in Quail et al. These appear to be the most recent and detailed numbers for regional beef packing concentrations and market shares (table 1). We use them as baseline estimates, $\hat{H}_k$s and $\hat{Q}_k$s, of the H_ks and Q_ks. For estimates of the elasticities of the national derived demand and supply curves for cattle, we adopt the figures $\hat{\eta} = -.53$ and $\hat{e} = 1.68$ obtained by Schroeter (1988) using annual data.

Baseline case estimates of the national average oligopsony distortion, $\hat{D}$, and the regional conduct parameters, $\hat{\alpha}_k$s, are also required. The test cases are designed to address

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[5] The expression of D^* in equation (7) embodies the assumption that equilibrium regional market shares in the test case will not be significantly different from those for the baseline case.

[6] Projecting the impact on quantities and prices in individual regions is a trickier matter not undertaken here. It clearly would be incorrect to simply move back along stationary regional supply curves until points were reached at which the estimated distortions, the D_k^*s, just "fit" between perfectly elastic demand curves and the supply curves. This would ignore the interdependence among regional supply curves as well as output price effects. By limiting attention to national quantity and (average) price, we avoid having to model the shifts in regional supply curves caused by price changes in adjoining regions.

[7] This technique is patterned after a similar exercise in Dickson and Yu.

Table 1.   Cattle Procurement Regions, Concentration Indices, and Slaughter Volumes for 1986

Region (k)	Geographical Description[a]	Herfindahl Index[b] (H_k)	Slaughter Volume (1,000 head)[b] (Q_k)
1	All of Washington, Oregon, Idaho, and Montana west of the Continental Divide	.5086	1,328
2	Northern California and Reno, Nevada	.3369	363
3	Southern California and Arizona	.1505	821
4	New Mexico and southwest Texas	.7718	23
5	All of Colorado, western Nebraska, and the southeastern corner of Wyoming	.3044	1,842
6	North and South Dakota and Montana east of the Continental Divide	.4620	438
7	Wisconsin, most of Minnesota, and northern Illinois	.3063	1,647
8	Iowa, eastern Nebraska, southern Minnesota, and Rockport, Missouri	.2068	6,386
9	Kansas (excluding the southwest corner), the western half of Missouri, and a northern slice of Oklahoma	.7277	1,087
10	Southwest corner of Kansas, the Oklahoma panhandle, the Texas panhandle, and three counties (Curry, Quay, and Union) in New Mexico	.2143	9,999
11	Remainder of Texas (excluding those parts in regions 4 and 10) and Oklahoma (excluding the parts in regions 9 and 10)	.1685	665
12	Eastern Missouri and southern Illinois	.3241	67
13	Indiana, Michigan, and Ohio	.0893	456

[a] These are rough geographical descriptions. Readers interested in the precise boundaries of the regions should refer to Quail et al.
[b] These data were provided by Bruce Marion.

current concerns about growing concentration or worsening oligopsonistic conduct. That is, they are characterized by $\alpha_k^* > \hat{\alpha}_k$ and $H_k^* > \hat{H}_k$. We deliberately bias our results in the direction of large apparent impacts by assuming current conduct is described by the noncooperative extreme, $\hat{\alpha}_k = 0$ for all k.

D represents a weighted average of regional values of $(p - c_k - w_k)/w_k$. Our estimate of it is based on industry financial data. Assuming that marginal (economic) costs are well approximated by average (accounting) costs, D can be estimated as before-tax earnings [revenue minus the sum of processing (or "operating") costs and livestock costs] as a proportion of livestock costs. The American Meat Institute (AMI) conducts annual financial surveys of meat packing firms in the U.S. Its *Annual Financial Reviews of the Meat Packing Industry* for 1979 through 1986 report figures for beef packers' operating costs that average 11.12% of sales and for livestock costs that average 87.62% of sales. Our baseline estimate of the distortion, therefore, is $(100 - 11.12 - 87.62)/87.62$ or 1.44%. Schroeter (1988) estimated oligopsony distortions in beef packing econometrically without reference to industry financial data. The simple average of his distortion estimates for the years 1979 through

1983 (the last year of his sample) is 1.03%. The close correspondence between this figure and the present method's estimate of 1.44% is reassuring.

These values for $\hat{D}$, and for $\hat{\alpha}_k$, $\hat{H}_k$, and $\hat{Q}_k$ for $k = 1, 2, \ldots, 13$, are inserted into equation (5) to obtain an estimate of the representative regional supply elasticity that is consistent with the baseline case's embedded assumptions about the national average distortion and regional conduct, concentration, and market shares. The resulting estimate is $\hat{\epsilon} = 18.3$. Obviously, this figure is much larger than any plausible estimate of the national average supply elasticity. Regional supply responses, though, reflect not only production responses but also cattle producers' opportunities to shift supplies among regions, in the short term, in response to transport-cost-compensated price differentials. The large magnitude of $\hat{\epsilon}$ suggests that these opportunities are extensive.[8] Nonetheless, sensitivity of our results to the mag-

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[8] Studies of dynamic comovements among regional cattle prices provide indirect evidence of the extent of spatial arbitrage. Schroeder and Goodwin found that price shocks in the eastern Nebraska market (in our region 8) trigger contemporaneous (same week) price reactions in the Texas panhandle (in our region 10) of at least 70% of the initial shock. Moreover, the Texas price adjustment to the Nebraska price shock is completed in only two weeks.

nitude of regional supply elasticity will be discussed.

In any particular simulation, the test case's values for the α_k^*s and the H_k^*s are then combined with $\hat{\epsilon}$ in equation (6) to project regional distortions for the test case. These are averaged in equation (7) to obtain D^*, the projected national average distortion. Finally, $\hat{D}$ and D^* are combined with estimates of e and η in equations (10) and (11) to determine average price and quantity effects.

Table 2 presents the results of simulations for two different test cases. The last two columns of the table give projections of test case, or "future," values for cattle market quantity and price. Recall that baseline, or "current," values of these variables are normalized to 100 and 1, respectively. The first line of the table projects the effects of increasing concentration to the extent of raising each regional Herfindahl index to the level $H_4 = .7718$, the highest measured value in 1986, while preserving completely noncooperative (Cournot) behavior in all regions.[9] Reference to table 1 shows that the degree of consolidation necessary to bring the industry to this level of concentration would be extensive: in 1986, the four largest regions (regions 5, 7, 8, and 10 with a collective market share of nearly 80%) each had Herfindahl index values of less than half of H_4. Region 9 is the only region with a 1986 Herfindahl index value approaching that of region 4, but regions 9 and 4 together accounted for only 4.5% of slaughter. The projected quantity and price effects associated with this very significant change, 1.08% and .64% respectively, are relatively small, however.

Of course, it is natural to suspect that the result of increased concentration would be not merely a less competitive Cournot equilibrium but an outcome reflecting greater concentration and a greater degree of oligopsonistic coordination. The second line of table 2 simulates the effects of a transition from Cournot conduct in all regions to pure monopsony in all regions. This represents the perfectly collusive limiting case of market coordination. When all α_k^*s = 1, concentration, measured by the H_k^*s, is irrelevant. Even for this extreme scenario, the magnitudes of the changes seem modest: quantity falls by 1.55% and price falls by .93%.

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[9] Herfindahl indices, sums of squared market shares, are often reported as whole numbers. We write them as decimals because this is how they enter our equations.

Table 2. Simulation Results

	Baseline Case Parameters			National Market Supply Elasticity ($\hat{e}$)	Derived Demand Elasticity ($\hat{\eta}$)	Test Case Parameters		Test Case Projections		
Distortion ($\hat{D}$)	Conduct ($\hat{\alpha}_k$)	Concentration ($\hat{H}_k$)				Conduct (α_k^*)	Concentration (H_k^*)	Distortion (D^*)	Quantity[a] (Q^*)	Price[b] w^*
1. .0144	all $\hat{\alpha}_k = 0$	all $\hat{H}_k$ = 1986 values	1.68	−.53	all $\alpha_k^* = 0$	all $H_k^* = .7718$	.0421	98.92	.9936
2. .0144	all $\hat{\alpha}_k = 0$	all $\hat{H}_k$ = 1986 values	1.68	−.53	all $\alpha_k^* = 1$	irrelevant	.0545	98.45	.9907
3. .0779	all $\hat{\alpha}_k = 0$	all $\hat{H}_k$ = 1986 values	3.36	−1.06	all $\alpha_k^* = 0$	all $H_k^* = .7718$	.2276	90.05	.9693
4. .0779	all $\hat{\alpha}_k = 0$	all $\hat{H}_k$ = 1986 values	3.36	−1.06	all $\alpha_k^* = 1$	irrelevant	.2949	86.26	.9570
5. .0779	all $\hat{\alpha}_k = 0$	all $\hat{H}_k$ = 1986 values	.84	−1.06	all $\alpha_k^* = 0$	all $H_k^* = .7718$	.2276	94.09	.9300
6. .0779	all $\hat{\alpha}_k = 0$	all $\hat{H}_k$ = 1986 values	.84	−1.06	all $\alpha_k^* = 1$	irrelevant	.2949	91.76	.9027

[a] Baseline case value, $\hat{Q} = 100$.
[b] Baseline case value, $\hat{w} = 1$.

These results are based on the estimates $\hat{D}$ = .0144, $\hat{e}$ = 1.68, and $\hat{\eta}$ = −.53, all of which are subject to challenge. We briefly consider the effects of changes in these values. Equations (5), (6), and (7) show that, with other parameters of the baseline and test cases held fixed, an increase in $\hat{D}$ will lead to an equiproportional increase in D^*. But since $D^* > \hat{D}$, the ratio $(1 + D^*)/(1 + \hat{D})$ will increase and, since the exponents in equations (10) and (11) are negative, Q^* and w^* will fall; that is, the projected quantity and price effects, $100 - Q^*$ and $1 - w^*$, will increase. Intuitively, a greater value of $\hat{D}$ can be reconciled with given baseline conduct assumptions only through a reduction in the estimate of the regional supply elasticity. The test case's increase in concentration or coordination will be more distortionary when coupled with the new lower elasticity estimate.

Differentiating equations (10) and (11) with respect to e and η for given D^* and $\hat{D}$, yields $dQ^*/de < 0$, $dQ^*/d\eta$, dw^*/de, and $dw^*/d\eta > 0$. Thus, increasing $\hat{e}$, other things being equal, will decrease Q^* and increase w^*; that is, increase the projected quantity effect and decrease the projected price effect. These results can be seen in figure 1 by visualizing the effects of a clockwise rotation of the supply curve about the point $(\hat{Q}, \hat{w})$. As supply becomes more elastic, preserving the same test case distortion, D^*, requires that Q^* fall and w^* rise. Similarly, increasing $\hat{\eta}$ in absolute value decreases Q^* and w^* thus increasing both quantity and price effect projections. This can be seen by visualizing a counter-clockwise rotation of figure 1's demand curve about the point $(\hat{Q}, 1 + \hat{D})$. As demand becomes more elastic, preserving a given value of D^* for the test case distortion requires that Q^* and w^* both fall. The observations of this and the previous paragraph imply that "large" projected quantity impacts will obtain with "high" values of $\hat{e}$, $|\hat{\eta}|$, and $\hat{D}$, while "large" projected price impacts will obtain with a "low" value of $\hat{e}$ and "high" values of $|\hat{\eta}|$ and $\hat{D}$.

Lines 3 and 4 in table 2 report the results of simulations for the two test cases considered above but with different baseline values for $\hat{D}$, $\hat{e}$, and $|\hat{\eta}|$. Values of $\hat{e}$ and $|\hat{\eta}|$ were set at twice our original estimates. Any error in our estimate of D would most likely have entered through the associated estimate of marginal processing cost: 11.12% of average revenue. Halving this figure leads to the estimate, $\hat{D}$ =

$(100 - 5.56 - 87.62)/87.62 = 7.79\%$.[10] This provides the value of $\hat{D}$ used in lines 3 and 4 of table 2. These choices for $\hat{e}$, $|\hat{\eta}|$, and $\hat{D}$, favoring a large quantity effect, do generate the prediction of a 13.74% decline in volume as the result of a transition to pure regional monopsony.

Lines 5 and 6 of the table pertain to the same test cases and, once again, the values for $\hat{D}$ and $|\hat{\eta}|$ used in lines 3 and 4. This time, however, the estimate of $\hat{e}$ is one-half of its original level. These choices, favoring a large price impact, achieve effects on price as large as 9.73% in the pure monopsony case.

Discussion and Conclusions

It is important, at this point, to reiterate the assumptions upon which our projections are based. First, the entire analysis takes static profit maximization as the maintained hypothesis and this, to be sure, is a theoretical limitation of our approach. If the dominant motivation of packing firms were some other goal, for example, profit-constrained market share maximization or any type of intertemporal objective, our results would be suspect. Second, all of the parameters characterizing the model's baseline case must be estimated and all of our estimation procedures (except perhaps the use of the Marion data for the H_ks and the $\hat{Q}_k$s) are subject to challenge. For this reason, we investigated the sensitivity of the procedure to alternative choices for $\hat{D}$, $\hat{e}$, and $\hat{\eta}$.

How do our results compare with those obtained using other methods? The three most often cited econometric studies of the effects of packer concentration on regional cattle procurement are Ward (1981); Menkhaus, St. Clair, and Ahmaddaud; and Quail et al. As Connor notes in his summary of these studies, they are in general agreement with respect to magnitudes of price effects. Each of them finds a price range between the samples' least and most concentrated market areas/time periods of 1.2% to 2.5% of the price level. In the simulations reflecting our judgment about "most plausible" parameter values (lines 1 and 2 of table 2), price effects are less than 1%. Note that the Ward/Menkhaus, St. Clair, and Ahmaddaud/Quail et al. price effects represent differences between the least and the most concentrated market areas present in their sam-

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[10] The corresponding estimate of regional supply elasticity, $\hat{e}$, is 3.39.

ples. It is highly unlikely that firm conduct across these areas actually spanned the range delineated by the baseline and test cases of the simulation reported in line 2 of table 2: from completely noncooperative to perfectly monopsonistic. Yet our projected price effects are smaller than their measurements. Therefore, our results indicate less danger of falling cattle prices, as a result of increased packer concentration or coordination, than do those from conventional econometric studies.[11]

The fact that our method produces small estimates of price impacts is primarily attributable to our high estimate of regional supply elasticity. In order to reconcile reasonable estimates of the current distortion with even completely noncooperative current conduct, a very high value of the elasticity of regional supply is required. High supply elasticities limit packers' abilities to benefit from concentration or input market coordination. Just as monopoly behavior becomes decreasingly "monopoly like" as demand elasticity increases, very high supply elasticities make even pure monopsony conduct relatively ineffectual, at least with respect to price effects.

It should be noted, however, that even quite small price effects can have significant effects on packer and feeder profit. AMI figures for the years 1979–86 report that livestock costs were nearly 88%, and before-tax earnings only about 1.25%, of total beef packing sales during this period. Thus a fall in cattle prices of only .5% has the potential to increase packers' profit by about 35%. Iowa Cooperative Extension Service figures (Futrell) estimate the returns from finishing yearling steers to Choice slaughter grade to have been approximately $49/head in the fourth quarter of 1990. The same .5% decline amounts to a $4.40/head profit loss on a 1,100 pound steer at $80/cwt. In other words, feeders' profit would decrease by nearly 9%.

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[11] The results of our sensitivity analysis do qualify this conclusion somewhat. As lines 5 and 6 of table 2 demonstrate, large price effect projections can be generated using parameter estimates significantly different from our "most plausible" values.

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