[*Jpn. J. Rural Econ.* Vol. 9, pp. 1-14, 2007]

Dynamics of Slaughter Weight Response to Market Price Changes for Japanese Beef Cattle

Atsushi Chitose* and Robert D.Weaver†

The objective of this paper is to investigate the dynamics of the slaughter weight of Japanese beef cattle. We extend the theory of and provide empirical results for the dynamics of slaughter weight. Our theoretical analysis shows that optimal slaughter weight increases with an increase in the price of either feeder calves or the flow output while it decreases with an increase in either feed price or interest rate. Slaughter weight response to beef price depends on the sex of the animal. An increase in beef price is found to reduce slaughter weight of steers (male animals with no flow output); however, the result for slaughter weight of female animals is ambiguous. Our time series analysis for the Japanese beef sector (1975-1999) finds evidence of long-run relationships between the average slaughter weight and prices (except for interest rate), consistent with theoretical response patterns. However, the empirical results indicate that no specific relationships exist between the average beef slaughter weight and prices in the short-run.

Key words: Japanese beef market behavior, optimal slaughter weight responses, dynamic optimization problem, time series analysis, unit root tests with seasonality and structural break, vector error correction model, long-run and short-run relationships.

1. Introduction

Total beef supply in the market is measured by total carcass weight from all cattle slaughtered, which can be decomposed into two components: the number of cattle slaughtered and average slaughter weight per head.[1)] However, the existing literature tends to focus on the first component as the source of beef supply dynamics. The mechanism of the adjustment in herd size (breeding stock) has been extensively explored both theoretically and empirically; see e.g., Foster and Burt [6]; Rosen [29]; Rosen, Murphy and Scheinkman [30]; Rucker, Burt and LaFrance [31]; and Schmitz [32]. The emphasis on breeding stock seems legitimate because total beef supply is determined largely by slaughter numbers while variances in slaughter weight are relatively small across slaughtered cattle. However, slaughter weight may be one of the producer choices particularly when the expansion of herd size is limited by the availability of fixed factors (e.g., land) or financial constraints. Given changes in market conditions, a beef producer would adjust the length of the feeding period and slaughter weight in order to maximize profit. More important is the fact that slaughter weight has actually increased, contributing to the increase in beef supply over a period of time. Such empirical evidence is striking in Japan. Average slaughter weight (in terms of carcass weight) increased by 28% for *Wagyu* steers and by 40% for dairy steers during the period

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*Tokyo University of Agriculture and Technology
†The Pennsylvania State University, USA

The authors are grateful to two anonymous referees of this journal for their valuable comments and suggestions. The first author also would like to thank Professors Yujiro Fuchino, Masaru Kagatsume and Yoshimi Kuroda for their advice and help during the early stages of this research. The earlier draft of this paper was written when the first author was on leave as a visiting scholar at Pennsylvania State University in 2003. He greatly acknowledges a hospitable research environment provided by Department of Agricultural Ecnomics and Rural Sociology, Penn State as well as financial support from Tokai University.

1976-1999.[2]

The objective of this paper is to investigate the dynamics of slaughter weight of Japanese beef cattle. Exploring the determinants of increase in slaughter weight is expected to contribute to a better understanding about the development of the Japanese beef sector. So far, less effort has been made to examine the mechanism underlying the increase in slaughter weight in association with the increase in feeding duration over time in Japan. The previous empirical studies are more likely to investigate the Japanese beef market structure from a holistic point of view, with less emphasis on changes in slaughter weight and feeding period; see e.g., Hotta [10], Matsubara [18], and Monma [22].[3] Furthermore, although the increase in feeding duration has been identified as one of the managerial problems by researchers specializing in beef production management, their analyses of causes of such a problem have been limited to somewhat speculative assertions (e.g., Kurihara [17], Miyazaki [21]). For example, Kurihara [17, pp. 24-26] argues that increases in slaughter weight and feeding period observed over the period 1960-1984 were attributed to increases in feeder calf price and in beef price relative to feed price. Yet, his discussion relies on a simple trend analysis, and is not based on a theoretically rigorous model. A lack of micro foundation is also acknowledged as a common weakness in the previous econometric studies on the Japanese beef supply. Their estimation models are likely to be built on the assumption that the supply dynamics can be explained by the cobweb theory in relation to biological characteristics instinct in beef production. In this respect our approach differs from the previous studies because the theoretical analysis explicitly plays a complementary role in conducting the empirical analysis.

We develop a theoretical model that is specified as a beef producer optimization problem in a dynamic setting and examine empirically the relationships between slaughter weight and market conditions using advanced time-series analysis methods. Our theoretical model is based on the classic Jarvis [13]-Paarsch [25, 26] line of livestock optimization models where cattle are viewed as both consumption and capital goods, with the time of slaughter and the feeding program (i.e., feed ration) as producer choice variables. We extend the Paarsch's dynamic model by including a flow harvest so as to generalize all this line of previous models. This generic model would illustrate the Japanese beef producer behaviors in relation to the three routes of supply dynamics in the beef market, noted by Morishima [24].[4]

Our empirical approach applied to time series data has several advantages in econometric specification over previous empirical studies on the Japanese beef supply. First, we exploit time series properties that are found through various tests including unit root tests and cointegration tests. The previous studies have paid less attention to time series properties. For example, time series data are assumed to be all stationary, and a fixed length of lags based on standard beef production practices is *a priori* specified in the beef supply response equation. Second, we attempt to distinguish the long-term and the short-term relationships to confirm empirical findings from the previous studies. This is because while as Kurihara [17] argues, some relationships seem to exist between slaughter weight (in association with a length of feeding period) and market conditions over time, other studies found such relationships vague in the short-run. For example, Inoue [11, pp. 72-73] found from the farm-level cross-section data that there is no correlation between total increase in cattle weight over a feeding period and feeder calf price. Chino [4, pp. 82-83] also found no correlation between a feeding period and beef price in the cross-section data. Unfortunately, theory is weak in this matter. Our theoretical model can posit only the qualitative direction of slaughter weight in response to price changes in the steady state which illustrates the long-run relations, with the short-run relations being left as an empirical issue.

The main results from our study are briefly provided. Our theoretical analysis shows that slaughter weight increases with an increase in the price of either feeder calf or flow harvest while it decreases with an increase in either feed price or interest rate. The results theoretically support Kurihara's assertion for the relationship between slaughter weight and feeder calf price. It is also shown that slaughter weight response to beef price depends on the sex of the animal. An increase in beef price is found to reduce slaughter weight of steers (male animals with no associated flow output); however, the result for slaughter weight of female animals is ambiguous. The econometric estimation found evidence of long-run relationships between average slaughter weight and prices (except for interest rates). Results are consistent with those for our theoretical analysis in response directions of slaughter weight to market price changes. However, it is also found that the impacts of prices on slaughter weight are small. Results for the short-run indicate that slaughter weight response to price changes is ambiguous in direction. This short-run result is consistent with the findings from the previous cross-section analyses by Chino [4] and Inoue [11].

In the next section, we present a dynamic optimization model that is applicable to various types of beef producers and show results for the rational adjustment in slaughter weight and time of slaughter by a beef producer to changes in market conditions. In the third section, the Japanese beef supply behavior is empirically investigated using monthly data that are disaggregated by beef cattle type for the period 1975-1999. Summary and conclusions are provided in the last section.

2. Theoretical Model

Our theoretical model is an extended version of Paarsh's dynamic model. Following Paarsh [25, 26], we specify an operator as facing an infinite horizon over which multiple placements (rotations) of animals are fed, and choosing a feeding program and duration of feeding to produce weight at slaughter and a flow product (e.g., calves) during the feeding period to optimize profits resulting from sales. To simplify, we assume: (*a*) homogenous age and quality (genetic properties) of cattle, (*b*) constant returns-to-scale; (*c*) constant prices for both inputs and outputs through a feeding period, (*d*) price-taking (i.e., exogenous price variables), and (*e*) a fixed, exogenous vector of price forecasts over the future.[5] Within this framework, we examine beef supply behavior in a model incorporating calf supply, animal weight at slaughter, feeding program, and time of slaughter. Then, we derive optimal slaughter behavior for specific animal types to motivate hypotheses of interest in our empirical study.

1) Model specification

The beef operation is specified as involving a weight gain and a flow output technology that is initiated at some time $\tau=0$ and is operated at each time t within a feeding period (rotation) of length T.

Weight gain for a typical animal is specified as:

$$\dot{w}(t) = h(w(t), f(t)),^{6)} \tag{1}$$

where $w(t)$ is the live weight of an animal and $f(t)$ the quantity of a bundle of variable inputs at time t. Because feed costs dominate variable costs in cattle feeding, hereafter we interpret $f(t)$ simply as feed. We assume $h(t)$ satisfies typical monotonicity and concavity properties: $h_w<0$, $h_f>0$, and $h_{ff}h_{ww}-h_{wf}^2>0$. The production function of the flow output, $y(t)$ is specified as

$$y(t) = g(w(t), f(t)),^{7)} \tag{2}$$

where $g(t)$ is assumed to satisfy $g_i>0$, $g_{ii}<0$ for $i=w$, f and $g_{ww}g_{ff}-g_{wf}^2>0$. For a cow-calf operation, $g(t)$ may be viewed as the probability function of calving. At the beginning of an arbitrary rotation at time $\tau=0$ we view the producer as choosing the feeding program $f(t)$ and feeding period length T for each of an infinite series of future rotations. At the end of each rotation, the producer is viewed as repeating this process. The present value of profit from holding an animal for a duration T is specified as

$$\pi(T)\int_0^T [qy(t)-cf(t)]e^{-rt}dt + pw(T)e^{-rT}-p^c-K \tag{3}$$

where q, c, and p respectively denote the unit price of a flow harvest, the unit price of feed, and the unit price of beef, while p^c is the price of a feeder calf, K denotes other fixed costs per animal, (e.g., sheltering, veterinary expenses, etc.), independent of animal age, and r is the instantaneous interest rate. The producer is assumed to maximize the present value of profits from animal feeding over perpetuity. With rotation number defined as an integer j, the producer chooses the feeding program $f(t)$ and the feeding period length T for each of an infinite number of discrete rotations as follows:

$$V^* \equiv V(f^*(t), T^*) = Max_{\langle f(t), T\rangle} V = \sum_{j=0}^{\infty} \pi(T) e^{-rTj} = \pi(T)(1-e^{-rT})^{-1} \quad (4)$$

where * indicates optimal level. The associated current-value Hamiltonian is defined at an arbitrary time t within a rotation as:

$$H(t) = qg(w(t), f(t)) - cf(t) + \phi(t) h(w(t), f(t)) \quad (5)$$

where $\phi(t)$ is a co-state variable and is interpreted as a current shadow value of the marginal unit of weight of an animal at time t. At each point in time t, necessary conditions for an optimal feeding program and weight trajectory are given by the following Euler conditions:

$$c = qg_f(t) + \phi(t) h_f(t) \quad (6a)$$

$$qg_w(t) + \dot{\phi} = (r - h_w(t))\phi(t) \quad (6b)$$

Equation (6a) defines an optimal policy rule for feeding at each point in time during a single rotation. The optimal feeding rule requires feed to be adjusted to set the marginal cost of feeding equal to the marginal revenue product from feeding an animal. Marginal revenue from feeding involves the sum of marginal revenue from the flow output and the increase in weight gain associated with a marginal unit of feed. Equation (6b) is interpretable as an intertemporal arbitrage condition that requires the marginal benefit from an increase in weight to equal marginal cost that consists of the shadow cost of interest $r\phi$ and an increase in the depreciation rate, $-\phi h_w$; see Paarsch [25].

Choice of slaughter weight and time of slaughter follows from the following necessary conditions that are derived as transversality conditions.

$$-\phi(T) + p = 0 \quad (7a)$$

$$qg(w(T), f(T)) + \phi(T) h(w(T), f(T)) - cf(T) - rpw(T) - rV^*(T) = 0 \quad (7b)$$

given the optimal feeding regime $f^*(t)$ that satisfies equations (6) in each single rotation.

Equation (7a) shows that feeding for weight gain continues until its shadow value equals its market value. Equation (7b) requires holding to continue until the marginal benefit from holding the asset (animal) equals the marginal cost of its replacement

Table 1. Comparisons of various models in optimal slaughter response directions

Model specification	Static single-rotation		Static perpetual-rotation		Dynamic perpetual-rotation	
Major restrictions on equation (7b)	$rV^*=0$, $f(t)=\bar{f}$ $g(t)=0$		$f(t)=\bar{f}$ $g(t)=0$		$g(t)=0$	
Sex Original model	Male Jarvis	Female[1)] Yver	Male Paarsch	Female[1)] Paarsch	Male Paarsch	Female Ours
$\partial T^*/\partial p \left[\partial w_T^*/\partial p\right]$	+	?	−	?	−[−]	?[?] ⟨−[−]⟩*
$\partial T^*/\partial p^c \left[\partial w_T^*/\partial p^c\right]$	0		+		+[+]	+[+]
$\partial T^*/\partial K \left[\partial w_T^*/\partial K\right]$	0	0	+	+	+[+]	+[+]
$\partial T^*/\partial c \left[\partial w_T^*/\partial c\right]$	−	−	0	0	?[−]	?[−]
$\partial T^*/\partial r \left[\partial w_T^*/\partial r\right]$	−	−	−	?[2)]	?[−]	?[−]
$\partial T^*/\partial q \left[\partial w_T^*/\partial q\right]$		+		?		?[+]

Note: 1) In these original models that are specified strictly for cow-calf operators, a calf is assumed to be self-supplied such that p^c is set equal to zero. 2) Paarsch [26, p. 645] shows that this sign is negative but it is incorrect. He ignores that the probability of calving which is specified as $g(t)$ in our model is time-varying.

* The signs in ⟨ ⟩ apply in case $\int_0^T (R^*(t) - R^*(T)) e^{-rt} dt < (p^c + K)$, where $R^*(t) = qg^*(t) - cf^*(t)$.

with a new asset (calf).[8] This equation is interpretable as a generic form of optimal policy for the time of slaughter because it generalizes the optimal conditions for cattle slaughter decisions available in the literature. That is, both Jarvis' [13] and Yver's [34] conditions for a single feeding rotation follow from restricting the opportunity cost of replacement (rV^*) to zero. Paarsch's [25, 26] case where no flow harvest is considered follows by setting $qg(T)=0$.

2) Optimal slaughter response to price changes

We derived optimal slaughter response to price changes by applying the method developed by Arnott, Davidson and Pines [1] and Paarsch [25]. Because this derivation requires somewhat complex mathematical manipulations including the phase diagram analysis, here only results are presented.[9] Table 1 shows the comparisons of various models (including our model in the last column) in optimal slaughter responses to price changes.

Comparing the last two columns indicates that slaughter weight would respond to changes in feeder calf price, feed price or interest rate in the same direction for both male and female animals in dynamic settings where slaughter weight is explicitly specified as one of the producer choice variables. Slaughter weight increases with an increase in feeder calf price, while it decreases with an increase in either feed price or interest rate. However, slaughter weight response to beef price depends on the sex of the animal. The theoretically derived relationship between slaughter weight and feeder calf price is consistent with Kurihara's assertion. This is because an increase in calf price would reduce the opportunity cost of continued feeding (rV^*), providing the producer with an incentive to extend the feeding period of the current rotation. However, our results cannot provide theoretical justification for his other assertion that an increase in slaughter weight was caused by an increase in beef price relative to feed price over time. While a decrease in feed price is found to increase slaughter weight, an increase in beef price is likely to reduce slaughter weight of steers (male animals with no flow output) and slaughter weight response for cows is indeterminate in sign.[10] This implies that the direction of slaughter weight response to the relative beef price depends on the magnitude of an impact of beef price on slaughter weight relative to that of feed price.

3. Empirical Analysis

In this section, we empirically investigate the response of average slaughter weight to changes in the market environment for Japanese beef during the period 1975-1999. The empirical analysis is focused on three types of beef cattle: *Wagyu* steers, female *Wagyu* and dairy steers fed for beef. We choose to analyze the monthly aggregate Japanese beef supply to allow identification of transitory response within feeding rotations as well as permanent response. The period of analysis is set from April 1975 to December 1999 (297 observations) to avoid an effect of mad cow disease (BSE) on domestic beef supply behavior.

We use average slaughter weight as a measure of w_T and nominal prices (p, c, p^c, q, K, r). Average slaughter weight was computed by dividing aggregate (carcass) weight of slaughter by aggregate slaughter number. The data on p and p^c are, respectively, the unit price of beef at slaughter (yen per kg) and the price per calf (yen per head) paid by feedlot farmers. In each case, these variables vary by animal type. In contrast, the data on the price of concentrate feed for beef cattle (c) and the interest rate of a long-term government bond (r) do not vary across animal type. Assuming that the male to female ratio for offspring is 0.5, we use the average price of male and female calves sold by cow-calf farmers as a measure of q, which is relevant only for female *Wagyu*. Note that the data for female *Wagyu* reflects both heifers fed for beef, as well as cows culled by cow-calf farmers and fattened for a short period before slaughter. Available data do not allow us to decompose data for *Wagyu* female animals into heifers and cows over the sample period. We assume that the animal specific supply functions are homogenous of degree one in prices and normalize all price variables using the aggregate price index of farm inputs (interpreted as a measure of K). The data were obtained from various issues of the Japanese official statistical publications: w_T from *Annual Statistics of Meat Marketing*

[19], all price variables except for r from *Survey Report on Prices and Wages in Rural Villages* [20]; and r from *Japan Statistical Yearbook* [12].

The theory presented above motivates the conceptual structure of the estimation models as well as hypotheses of interest. In particular, the theory motivates consideration of reduced form slaughter weight supply functions. Empirically, consistent with separability of weight gain functions across animal types, we proceed under the maintained hypothesis that these functions are independent. This assumption is consistent with independence of animal type specific operations observed in Japan. Given that our data measure average weight at the sector level, we cannot rule out endogeneity of prices and short- or long-run structural relationships across prices and average weight that go beyond those motivated by our supply theory. Thus, based on our finding of cointegration among the variables, we adopt a vector error correction model (VECM) as an empirical approach that is relatively free of *a priori* restrictions on causal structure, yet allows investigation of both long-run and short-run behaviors.

Consistent with our maintained hypothesis that the weight gain technologies are separable across animal types, we exclude the price of calves, q from the *Wagyu* and dairy steers models. We estimate three animal type specific VECMs (subsystems): (*i*) *Wagyu* steers and (*ii*) dairy steers, as determined by the vector (w_T, p, p^c, c, r), and (*iii*) female *Wagyu* as determined by (w_T, p, p^c, c, q, r). The interest rate (r) is specified as an exogenous variable in each subsystem because the long-term government bond interest rate (r) evolves outside the beef sector. Initial inspection motivated logarithmic transformation of the data to linearize variation and relationships.

Estimation proceeded in three steps. In the first step, univariate stationary properties of each variable were investigated using (*i*) the augmented Dickey-Fuller (ADF) test, (*ii*) the seasonal unit test (Beaulieu and Miron [2]) and (*iii*) the unit root test with a structural break (Zivot and Andrews [35]). Because most variables (except for c and r) in our sample exhibit distinct seasonal variations, whether such seasonality is attributed to deterministic or stochastic properties was examined following Beaulieu and Miron. The Zivot and Andrews unit root tests were pursued because structural change within the sample could not be ruled out *a priori* and ADF tests are more likely to reject a null hypothesis of $I(1)$ when structural change takes place within a sample period. The possibility of structural change in the Japanese beef sector is highlighted by policy changes such as beef import liberalization in the early 1990s that have affected the sector during the sample period.

In the second step, we examined evidence of long-run relationships in each animal type subsystem following the Johansen procedure; see e. g., Johansen and Juselius [14]. After the cointegration rank was determined, hypotheses tests were implemented to determine appropriate specifications of VECMs. In the last step, the transitory response behavior of the series was examined using impulse response functions and forecast error decomposition based on the estimated structural VECMs.

Table 2 presents a summary of ADF and the seasonal unit tests. In these tests, deterministic components in the ADF regression for each series were chosen following Dolado, Jenkins and Sosvilla-Rivero [5], and the appropriate lags were identified relying on Akaike's information criterion (AIC), Schwartz information criterion (SIC) and F-tests. The ADF tests indicated each series is $I(1)$. Although these results are not reported here, we performed the ADF tests with a null of $I(2)$, and found the null could be rejected for each series. These results are supported by unit root tests with a structural break; see Appendix.[11] However, seasonal unit root tests provided evidence that, in some cases, was inconsistent with these inferences. In particular, most variables for the female *Wagyu* subsystem are found to be better approximated by $I(0)$ at zero frequency, implying stationarity holds for levels. No stochastic seasonality was found. The existence of a unit root was rejected at each seasonal frequency for each series and for each animal type, except for the beef price of *Wagyu* females. The absence of seasonal unit roots for most series suggests that no long-run relationships exist at any seasonal frequency, for

Table 2. Results of unit root tests

	ADF tests[a)]			Seasonal unit root tests[b)]							
	Deterministic components				0	π	$\pi/2$	$2\pi/3$	$\pi/3$	$5\pi/6$	$\pi/6$
		Lags	t-stat.	Lags	π_1	π_2	$F_{3,4}$	$F_{5,6}$	$F_{7,8}$	$F_{9,10}$	$F_{11,12}$
					Wagyu steers						
w_T	int, sd, trend[c)]	19	−1.84	2	−4.36	−4.77	17.57	23.21	20.12	18.91	8.62
p	int, sd	8	−2.06	0	−2.35	−5.92	25.91	35.57	31.14	27.42	27.53
p^c	int, sd, trend	8	−2.47	0	−2.99	−4.17	29.79	32.29	31.85	40.06	21.66
					Dairy steers						
w_T	int, sd, trend	2	−1.54	1	−1.45	−4.77	22.34	26.43	29.54	24.99	34.86
p	int, sd, trend	7	−2.50	4	−1.80	−4.06	16.07	27.90	13.73	19.86	26.67
p^c	int, sd, trend	15	−1.53	4	−1.53	−3.94	13.92	22.15	15.93	36.18	12.28
					Wagyu females						
w_T	int, sd, trend	2	−2.22	1	−3.67	−3.64	21.45	18.89	26.80	16.32	28.08
p	int, sd	12	−2.78	1	−2.78	−1.61	2.46	5.28	7.38	2.01	16.02
p^c	int, sd	8	−2.62	0	−2.87	−4.08	28.65	25.69	36.13	34.67	30.87
q	int, sd	9	−2.57	1	−3.29	−5.12	20.10	21.78	25.15	18.99	22.96
c	int, trend	6	−2.50	1	−3.28	−4.23	16.43	12.95	18.31	21.93	19.78
r	int, trend	4	−1.80	0	−1.54	−4.36	28.08	33.46	18.40	28.90	22.80

Note: a) Critical values at the 5% level are: −2.88 for the ADF regression with intercept/seasonal dummies; and −3.43 for intercept/seasonal dummies, and time trend (for the number of simulated observations equal to 250, Fuller [7, p.371]). b) Critical values at the 5% level for π_1, π_2, and F: π_{odd}, π_{even} are: −2.76, −2.76, 6.26 for the regression (int, sd); −3.32, −1.88 and 2.97 for the regression (int, trend); and −3.28, −2.75, and 6.23 for the regression (int, sd and trend). These are from Table A.1 (for the number of simulated observations equal to 240) in Beaulieu and Miron [2, pp.325-326]. c) int=intercept, sd=seasonal dummies.

any of the animal types. On this basis, we concluded that seasonal cointegration tests are not necessary in our analysis.[12)]

To consider multivariate relationships, first we perform Johansen cointegration rank tests. The cointegration rank tests are based on maximum likelihood estimation of an unrestricted autoregressive (VAR) model. However, the validity of statistical inferences from such estimation depends on the validity of the VAR specification including multivariate normality and time- invariant covariances. As Hendry and Juselius [9] note, these specifications deserve empirical verification. For our data, preliminary VAR estimation provided strong evidence of violation of the Gaussian assumption even for extended lag lengths. This problem was partially corrected after intervention dummies were included in each subsystem to eliminate apparent outliers in differenced series.[13)] Because some series were found to have nonzero means in differences, the VAR model is specified such that the cointegration space contains a linear trend. Rather than imposing *a priori* restrictions on the cointegration space, the presence of a linear trend in long-run relationships is empirically examined. Based on the Hannan-Quinn (HQ) criterion and the analysis of residuals, the lag order in VAR models was chosen to be two for *Wagyu* steers and four for both dairy steers and *Wagyu* females.[14)]

Cointegration rank indicates the number of cointegrating vectors that characterize multivariate relationships in a long-run perspective. Table 3 presents the results of cointegration rank tests. It is found that the cointegration rank for both *Wagyu* steers and dairy steers is two while the rank for female *Wagyu* is three.[15)] With the rank specified, various hypotheses tests were implemented. The null hypothesis of univariate stationari-

Table 3. Results of cointegration rank tests

		Wagyu steers		Dairy steers				*Wagyu* females		Critical values (90%)[b]		
r	p-r[a]	λ-max	trace	λ-max	trace	r	p-r	λ-max	trace	p-r	λ-max	trace
		—	—	—	—	0	5	70.25	167.89	5	23.72	82.68
0	4	58.86	99.55	44.01	90.24	1	4	54.91	97.64	4	19.88	58.96
1	3	24.22	40.69	27.09	46.23	2	3	25.76	42.73	3	16.13	39.08
2	2	10.01	16.46	13.16	19.14	3	2	9.70	16.97	2	12.39	22.95
3	1	6.45	6.45	5.97	5.97	4	1	7.27	7.27	1	10.56	10.56

Note: a) r (non-italic) is the number of cointegration ranks, and p (non-italic) is the number of variables. b) The critical values come from Hansen and Juselius [8, pp. 80-81]. However, these critical values should be treated with caution because the VAR model that contains several intervention dummies changes the asymptotic distribution of the trace statistic on which critical values presented here rely.

Table 4. Coefficients of cointegration vectors

	Wagyu steers[a]				Dairy steers[b]				*Wagyu* females[c]					
	Coeff.	S.E	Coeff.	S.E	Coeff.	S.E	Coeff.	S.E	Coeff.	S.E	Coeff.	S.E	Coeff.	S.E
w_T	1.000	—	—		1.000	—	—	—	1.000	—	—	—	—	—
p	0.306	0.046	1.000	—	0.832	0.036	1.000	—	0.053	0.040	—	—	1.000	—
p^c	−0.166	0.025	−0.405	0.051	−0.305	0.024	−0.309	0.031	−0.028	0.010	1.000	—	—	—
q	—	—	—	—	—	—	—	—	−0.028	0.010	−0.801	0.027	—	—
c	0.328	0.031	—	—	0.206	0.031	—	—	0.455	0.039	—	—	3.602	0.372
r	—	—	−0.149	0.037	—	—	—	—	—	—	—	—	—	—
trend	—	—	0.000	0.000	—	—	0.001	0.000	—	—	—	—	0.005	0.000

Note: a) The number of the identification restrictions imposed on this subsystem is 4, and its test (likelihood ratio) statistic is $\chi^2(2)=5.29$, with the critical value (5%) equal to 5.99. b) The number of the identification restrictions imposed on this subsystem is 5, and its test statistic is $\chi^2(3)=2.78$, with the critical value (5%) equal to 7.81. c) The number of the identification restrictions imposed on this subsystem is 12, and its test statistic is $\chi^2(6)=12.15$, with the critical value (5%) equal to 12.59. Note that coefficients for p and q are set to be equal in the first equation. The alternative model that appears to be more preferable but does not satisfy the rank condition has the following cointegrating vectors: $[1, 0.260, -0.084, -0.106, 0.295, 0, 0]'$, $[0, 0, 1, -0.787, 0, 0, 0]'$, $[0, 1, -4.70, 0, 0.843, 0, 0.02]'$. The test statistics for overidentification restrictions is $\chi^2(4)=6.64$, with the critical value equal to 7.81.

ty, $I(0)$, for each series was reexamined within the multivariate setting of the VECMs for each animal type. In each case, the hypothesis is tested by imposing restrictions on each cointegrating vector, restricting parameters of all other variables in the vector to zero while leaving all parameters in other cointegrating vectors unrestricted. Results indicated that the hypothesis of stationarity can be strongly rejected for each series, confirming the univariate results such that each series is well approximated as $I(1)$ within the multivariate analysis. For each animal type VECM, the null hypothesis of no linear trend in the cointegration space was rejected, as was the null of no interest rate in the cointegration space for dairy steers and *Wagyu* females but not rejected for *Wagyu* steers (at a 5% significance level).

The presence of multiple cointegration vectors suggests that there exist long-run relations across slaughter weight and the mechanisms generating the determinants of slaughter weight. This may reflect the fact that prices are likely to co-move with each other due to the presence of a common stochastic shock. These results may follow market efficiency and integration resulting from arbitrage as discussed by Chambers and Baily [3] or Pindyck and Rotemberg [28], among others. This would especially be expected for the relationship between paid and received calf prices (p^c and q) in the female *Wagyu* subsystem. It would be expected that these prices would be cointegrated if they were generated by a competitive calf market where arbitrage drove these prices toward a long-run relationship. Evidence is consistent with this type of efficiency in arbitrage between buyers and sellers (see the third and the forth columns for *Wagyu* females in Table 4).

Table 4 presents the estimated coefficients of the cointegration vectors for each animal type subsystem. The cointegrating relationships were identified using the overidentifying restrictions on the VECM to satisfy the rank condition (Johansen and Juselius [14, p. 15]). Here, the focus of our interpretation is placed on the relation existing in beef production. We interpret the first vector (the first column) in each subsystem in Table 4 to represent the long-run relations for Japanese beef production. The signs are consistent with those derived theoretically in the previous section, except for those for the optimal slaughter response behavior with respect to the interest rate (r). Recall theory suggested $\partial w_T/\partial p<0$ for steers, and for each animal type $\partial w_T/\partial p^c>0$, $\partial w_T/\partial c<0$, and $\partial w_T/\partial q>0$. Statistically significant evidence of long-run relations consistent with this theory was found for each of the three subsectors. These results support the conclusion that the theoretical model provides a useful characterization of the long-run Japanese beef cattle slaughter behavior. Estimated coefficients for the interest rate were not statistically different from zero. This result suggests Japanese beef producers are not sensitive to interest rate changes. Two plausible reasons are considered. First, an interest rate alone does not represent the opportunity cost of capital in beef production management. Mori [23, pp. 93-97] argues that beef producers may perceive not only interest rates but also transaction costs (i. e., agency costs) as costs incurred in borrowing money. Second, Japanese beef producers have preferential access to financial markets shaped primarily by institutional credit through agricultural cooperatives. The latter reason seems more probable since many beef producers have faced a debt problem that may have been attributed in part to the lack of farmers' financial discipline given favorable financial markets.

Although results are interpreted with some caution due to statistical properties found in our data (see note 15); the empirical results suggest that long-run price responsiveness of beef slaughter weight in Japan is small.[16] A comparison of coefficients across the three animal types suggests that slaughter weight w_T is most price-responsive for dairy steers and least responsive for *Wagyu* females. In addition, the speed of the adjustment to the long-run equilibrium is found to be slow. The estimated coefficients of the adjustment in w_T for the cointegration vector associated with beef production are -0.052 for *Wagyu* steers, -0.102 for dairy steers and -0.135 for female *Wagyu*. In each case, the estimated coefficients are statistically significant at a 1% significance level.

To examine supply response to short-run (transitory) changes, we rely on impulse response and forecast error decomposition.[17] Due to space limitation, only the main findings from such analyses are presented. The results indicate that the short-run price responsiveness of slaughter weight is small. Variance decompositions reveal the presence of strong inertia in the short-run dynamics of the slaughter weight response. The shock to w_T may be interpreted as resulting from unanticipated, random events such as accidents, injuries and technical efficiency or progress. In any case, this shock accounted for more than 90% of forecast error variance for average slaughter weight in a 10-month-ahead forecast. Specific results were 96% for *Wagyu* and dairy steers and 92% for female *Wagyu*. The analysis of impulse responses based on a 95% confidence band shows that slaughter response with respect to a transitory shock to any of the price variables is ambiguous in direction. This result holds for each type of beef animal. For example, either a negative or a positive transitory response of slaughter weight to a shock on feeder calf price is possible for *Wagyu* steers.

The empirical results show that although the response direction of average slaughter weight is consistent with the theory in the long-run, the impact of price on slaughter weight is small. Results for the short-run indicate beef slaughter weight response to price changes is small and even ambiguous in sign. These short-run results suggest the absence of specific relationships between slaughter weight and prices in the short-run, which is consistent with the results from the previous studies based on cross-section data (Chino [4] and Inoue [11]).

4. Summary and Conclusions

This paper has empirically analyzed the dynamic behavior of beef supply with a focus on the adjustment in slaughter weight in response to changes in economic environments. Our analytical approach differs from the previous studies on the Japanese beef supply with respect to the following points. First, we shed light on the mechanism underling the dynamics of slaughter weight that has received less attention in the previous studies despite the fact that the increase in slaughter weight has contributed to the increase in total beef supply over time. Second, the empirical analysis is explicitly associated with the theoretical analysis that motivates a conceptual structure of the estimation models as well as hypotheses of interest. Third, in the empirical analysis we exploit time series properties that are found through various tests; moreover, we can distinguish the long-term and the short-term relationships through vector error correction models.

Our theoretical analysis shows that optimal slaughter weight increases with an increase in feeder calf or flow output prices while it decreases with an increase in feed price or interest rate. Slaughter weight response to beef price depends on the sex of the animal. An increase in beef price is found, in general, to reduce slaughter weight of steers (male animals with no associated flow output); however, the adjustment direction for slaughter weight of female animals is ambiguous. Our theoretical model is viewed as a generalized model that nests within it previous models for cattle slaughter decisions available in the literature (e. g., Jarvis [13], Paarsch [25, 26], Yver[34]).

Our empirical analysis of time series for Japanese beef found evidence of long-run relationships between the average slaughter weight and prices (except for the interest rate), which are consistent with theoretically derived patterns. However, the impacts of prices on slaughter weight are small, in particular in the short-run. The empirical results suggest that adjustment in slaughter weight plays a role in the long-run as a component of beef supply adjustment. In the short-run, the role of slaughter weight is limited for the Japanese case, however. This result seems reasonable when the producer is viewed as choosing feeding practices on the grounds of the long-run profit. The finding that short-run response is small suggests that beef supply chains may not be exposed to short-run supply variation induced by response to unanticipated shocks to the market environment.

1) The conversion ratio from dressed weight to carcass weight also serves as one component of total beef supply (in carcass weight); however, its effect on change in total beef supply over time is considered negligible compared to those of the other two components.
2) Authors' computation based on data from *Annual Statistics of Meat Marketing*, various issues [19].
3) Hotta [10] has specified the feeding period to be exogenous in his econometric model. Although Matsubara [18] and Monma [22] have both treated slaughter weight as an endogenous variable in their econometric models, an equation for slaughter weight (in terms of carcass) is simply specified with the lagged terms of beef price and of slaughter weight as explanatory variables. In contrast to these studies, Komaki [16] has examined the Japanese beef supply structure with slaughter weight specified as a variable of beef supply. His econometric model using time series data is similar to our econometric models; however, the motivation of study differs between the two studies and our approach is more sophisticated with solid micro foundation and more prudent estimation procedures.
4) Morishima [24, pp. 6-11] notes that when the beef price falls, cow-calf operators reduce the number of cows while feedlot operators may sell steers earlier. This implies that the adjustment in feeding period length acts as a determinant of domestic beef supply for both cow-calf operations and feedlot operations.
5) Paarsch [25, 26] used similar assumptions. Assumption (*b*) implies that the producer's choice of the time and weight of slaughter is separable or recursive with respect to the choice of herd size. Assumption (*c*) suggests that the price of fresh beef is independent of the age of the animal at slaughter.
6) In a dynamic sense, this equation is an autonomous function since it is independent of time, or the age of an animal. Further, season-

ality within a single feeding rotation is ruled out, again to simplify and facilitate use of phase diagrams for the analysis of optimal response behavior in the steady state.

7) Two assumptions associated with $g(t)$ are worthy of note. First, it is specified as an autonomous function as assumed for $h(t)$. Second, a non-negativity condition is not explicitly specified to simplify the exposition of analysis. Because a female calf cannot generally produce flow harvests immediately after it is introduced into a herd, $g(t)$ must be equal to zero at the early stage of growth. If the non-negativity condition is explicitly specified in the model, there may be the corner solution for which an animal is sold before it yields a flow harvest. Third, $g(t)$ may be viewed as the production function of milk for dairy farming. This implies that the model can be applied to a dairy operator's optimization problem. However, in actuality dairy farmers are most unlikely to decide the time of culling cows, faced with changes in beef price in the market. Culled cows are viewed as by-product in dairy farm management. Thus, the analysis of slaughter responses for dairy cows is excluded from the present study.

8) The opportunity cost of replacement, rV is equivalent to Faustmann land rent in forestry management for which land is only a fixed resource necessary for forest growth (Paarsch [26, p. 643]). For cattle feeding, rV is interpreted as the imputed value of Ricadian rent to fixed resources, that is, human and physical capital used in perpetuity (returns/costs). Note that equation (7b) can be rewritten as $\phi(T)h(T)=r(1-e^{-rT})^{-1}[pw(T)-\int_0^T(R^*(t)-R^*(T))e^{-rt}dt-p^c-K]$, where $R^*(t)=qg^*(t)-cf^*(t)$. In this form, it is clear that (7b) is the classical condition for optimal asset replacement in an infinite time horizon operation shown by Perrin [27].

9) The derivation is available from the authors upon request.

10) The response direction to beef price for female animals depends on the sum of the calf price and other fixed cost relative to the sum of the discounted values of differences in instantaneous net revenues (i. e., $R^*(t)$ minus $R^*(T)$ at each point in time over the period, where $R^*(T)=qg^*(t)-f^*(t)$). This implies that if the feeding duration is relatively short with a small variation in instantaneous net revenues throughout the period, the direction of slaughter weight response would be the same as for male animals such that $\partial w_T^*/\partial p<0$ and $\partial T^*/\partial p<0$.

11) The breakpoint differs across variables in each subsystem. However, in many price series, structure change appears to have taken place sometime in the mid-1980s to the early 1990s, a period right before or around trade liberalization of beef imports in 1991. These results imply that trade liberalization had more impact on the Japanese beef industry during its transition period than after its implementation.

12) In addition to nonstationary properties, we attempted to identify other properties of each series that may distort statistical inferences in the subsequent multivariate analysis. Our analysis relied on standard statistics and graphical inspection for each series with due attention to the presence of outliers ($>\pm 3\hat{\sigma}$) in both levels and differences. Graphical inspection indicated that all outliers appear to be spikes in a single observation period. Results form the Zivot and Andrews [35] unit tests (the nonexistence of a unique breakpoint) ensure that we can treat outliers using intervention dummies in the VARs in the next step.

13) Dummy variables are included in the VAR model to ensure normality can be accepted. Univariate analysis suggests that outliers in each series may not have effects on the long-run relationships in any subsystem over the sample period. This implies that the cointegration tests with the presence of structural break in the deterministic trend (Johansen, Mosconi and Nielsen [15]) do not apply to our sample.

14) Although other information criteria such as Schwartz (SIC) and Akaike (AIC) were also considered, the HQ is used as a principal criterion following Johansen, Mosconi and Nielsen [15, p. 233].

15) These results must be interpreted with some caution. First, the vector of residuals from the VAR model used for the rank test does not satisfy the multivariate normality condition at the 5% level in the subsystem. Yet, we interpret the implications of this violation for statistical inference as not serious because excess kurtosis (fat-tailed distributions) was the likely cause of violation (for feed price (c)). Second, use of critical values in Table 3 must be based on recognition that the VAR model contains several intervention dummies that change the asymptotic distribution of the trace statistic on which critical values used here rely. These limitations were resolved by performing the careful graphical analysis (e. g., parameter constancy) and by implementing alternative VAR models with some modification in order to confirm the robustness of the presented results.

16) The coefficients (normalized by slaughter

weight) cannot be interpreted as elasticities which show the relative magnitudes of slaughter response with respect to prices *ceteris paribus*. Recall that no exogeneity hypothesis is imposed *a priori* so coefficient estimates reflect both direct and indirect effects of price change on slaughter weight, with the latter caused by interactions with all other variables in the long-run.

17) The restricted VECM identified in the cointegration analysis is used as a base for assessing short-run slaughter response. Transformation from a reduced form of VECM to a structural moving average (MA) representation requires further restrictions for identification. Restrictions are imposed on the leading coefficient matrix of the structural VAR, given several relevant assumptions. First, because our interest in the analysis is the slaughter response, we allow slaughter weight to be least restricted so that it can be responsive to contemporary prices. Second, feed price is not contemporarily affected by any variable in the beef sector. Because almost all livestock feeds are imported in Japan, the price of feed is more likely to be determined by the world price and foreign exchange rates as well as demand in the other livestock sectors (i.e., poultry, hogs and dairy). Third, beef price is affected to a limited extent by contemporary changes in supply factors (slaughter weight and calf price) because beef price is determined also by factors of consumer demand. Forth, the price of calves is affected by contemporary prices of beef and feed because they are both crucial determinants of the profitability of a calf that is viewed as a capital good. These assumptions are generally supported by statistical properties found in the sample such as partial correlations (Swanson and Granger [33]).

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(Received September 5, 2005 ; accepted September 6, 2006)

Appendix: Results of unit root tests with a structural break

	Model (A)[a, b, c]			Model (B)			Model (C)		
	Break point		t-stat	Break point		t-stat	Break point		t-stat
				Wagyu steers					
w_T	1986	1	−3.59	1991	1	−3.35	1987	11	−4.38
p	1985	10	−3.76	1990	1	−2.53	1985	10	−3.68
p^c	1985	6	−3.89	1989	1	−2.76	1985	6	−3.90
				Dairy steers					
w_T	1995	10	−2.75	1990	1	−3.40	1987	4	−4.24
p	1991	1	−3.98	1988	4	−3.69	1991	1	−4.32
p^c	1992	2	−3.13	1987	11	−3.06	1985	8	−3.28
				Wagyu females					
w_T	1975	10	−3.97	1976	11	−4.04	1976	7	−3.99
p	1985	7	−4.10	1990	2	−3.19	1985	7	−3.99
p^c	1986	2	−4.01	1989	4	−3.03	1986	2	−4.00
q	1985	8	−3.21	1988	10	−2.11	1985	4	−3.38
c	1985	9	−3.83	1976	2	−3.11	1985	9	−4.10
r	1996	7	−3.72	1992	4	−3.75	1989	11	−4.89

Note: a) Model (A) allows for a one-time change in the level of the series; Model (B) allows for a one-time change in the slope of the trend function of the series; and Model (C) combines changes in the level and the slope of the trend function (Zivot and Andrews [35]). b) Deterministic components and lags for each series follow those specified in its ADF egression (see Table 2). c) Critical values at the 5% level are: −4.80 for Model (A), −4.42 for Model (B), and −5.08 for Model (C), given infinite number of simulated observations, Zivot and Andrews [35, pp. 256-257].