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Estimating Structural Change in the Japanese Beef Import Market in the Wake of BSE: A Smooth Transition Approach

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We assess the effect of bovine spongiform encephalopathy (BSE) on Japanese beef imports from the United States and competing suppliers. Using a source-differentiated almost ideal demand system of beef imports with endogenous smooth transition functions, we find that a nonlinear structural change has occurred in the Japanese beef import market in the wake of BSE. The BSE outbreaks led an instantaneous, persistent impact on Japanese beef imports lasting over a decade, causing a significant shift in Japanese consumer preferences for beef imports from different origins. Over half of the estimated expenditure, own-price, and cross-price elasticities have changed in the aftermath of BSE, and some have not returned to their pre-BSE levels even after the trade recovery period.

Key words: beef imports, smooth transition function

Introduction

Prior to 2003, the United States was the world's largest exporter of beef and beef offal products, exceeding \$3.5 billion. Over 86% of U.S. beef products by volume were exported to Japan, South Korea, China, Canada, and Mexico. However, the discovery of bovine spongiform encephalopathy (BSE, or mad cow disease) in the state of Washington in December 2003 prompted an immediate ban on U.S. beef exports to nearly every primary destination market, causing severe losses to the U.S. beef industry. Coffey et al. (2005) estimated that the associated costs to the U.S. beef industry due to BSE for the year 2004 alone were \$200 million resulting from lower export sales and a reduction in unit prices. The U.S. Meat Export Federation estimated that the 10-year cumulative losses of U.S. beef trade as a result of the 2003 BSE outbreaks were \$16 billion, with most of the predicted losses occurring in the first 3 years. Peterson, Grant, and Sydow (2017) developed a global partial equilibrium simulation model of meat production and trade and found that U.S. beef exports would have been 2 million metric tons, or \$6.1 billion, higher if the BSE outbreaks had not occurred. While Mexico and Canada reopened their markets to U.S. beef relatively quickly, other markets in Asia remained closed for a much longer period. For example, Japan and South Korea suspended all imports of U.S. beef through 2005/2006, after which both countries began allowing imports of beef from the United States for cattle aged less than 21 and 30 months, respectively. China banned

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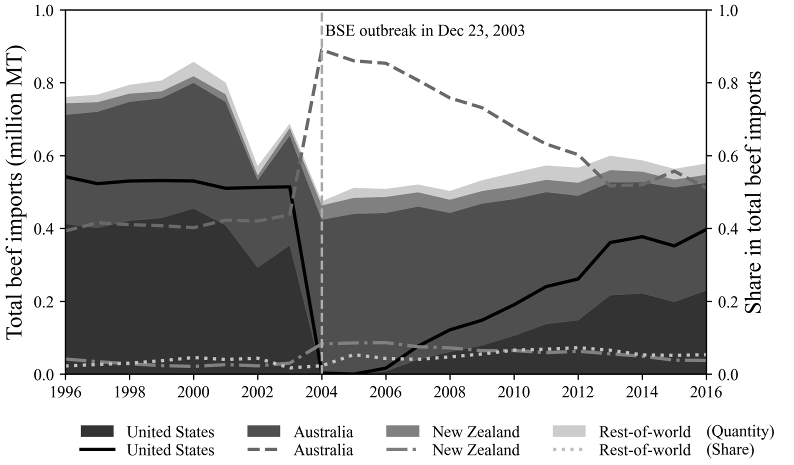

Figure 1. Japanese Total Beef Imports by Origin, 1996–2016

Notes: Total beef imports include products in the Harmonized System (HS) of commodity classification: 0201, 0202, 020610, 020621, 020622, 020629 and 160250.
Source: Authors' calculation using data from Global Trade Atlas.

imports of U.S. beef until September 2016, when China announced that it would begin allowing imports of U.S. beef from cattle aged less than 30 months, provided U.S. exporters complied with China's traceability and quarantine rules.

Before the discovery of BSE in the United States, over half of Japanese beef imports were from the United States, and Japan accounted for one-third of total U.S. beef exports. Figure 1 presents annual Japanese beef imports by exporting country from 1996 to 2016. In 2003, Japan imported 687,884 metric tons ($2.7 billion) of beef products. U.S. share of beef exports to Japan was 51% by volume, followed by Australia and New Zealand at 44% and 3%, respectively. However, Japan's total beef imports fell 31% in 2004, immediately following the U.S. BSE crisis. Australia and New Zealand, on the other hand, experienced notable increases in their beef exports to Japan: Australia's share of beef exports to Japan more than doubled in 2004, while New Zealand tripled its market share.

As of 2016, Japan's beef imports from the United States reached 230,000 metric tons, roughly 65% of its pre-BSE level in 2003.[1] Several factors may explain the slow recovery of U.S. beef exports to Japan. First, the BSE incident may have altered buyers' preference for beef imported from different suppliers in the Japanese market. Further, competing suppliers may have elevated their promotional strategies in the Japanese market in an attempt to increase market share. Second, although Japanese consumers reestablished confidence in demand for imported U.S. beef products, U.S. exporters have undergone changes in product compliance standards and marketing channels that could pose new hurdles for the recovery of U.S. beef exports. Third, the BSE events may have shifted Japanese purchases from imported to domestically produced beef products. For example, Japan's consumption of imported beef products had declined after 2003, while its consumption of domestically produced beef remained quite stable. At a minimum, this suggests that the U.S. BSE outbreaks affected Japan's beef import purchases more severely than its domestic beef purchases, leading to increased availability of substitute beef products from nontraditional suppliers. Thus, the absence of, and habit formation away from, U.S. beef may have resulted in a longer-term structural change in Japan's demand for beef imports.

Extensive research has been conducted to examine structural changes in meat demand (Chavas, 1983; Eales and Unnevehr, 1988, 1993; Moschini and Meilke, 1989; Choi and Sosin, 1990; Alston and Chalfant, 1991; McGuirk et al., 1995; Davis, 1997; Mangen and Burrell, 2001). Jin and Koo

[1] By value, Japan imported $1.49 billion of U.S. beef products in 2016, or 92.5% of its pre-BSE level in 2003.

(2003) and Ishida, Ishikawa, and Fukushige (2010) investigated the impacts of BSE and foot-and-mouth diseases on meat demand and found that animal disease outbreaks resulted in significant structural change in domestic demand for meat products. Peterson and Chen (2005) estimated a linear-switching Rotterdam model to assess the impact of BSE outbreaks on Japanese retail meat demand. Kawashima and Puspito Sari (2010) applied the Kalman filtering method to estimate a time-varying Armington model and found that BSE outbreaks resulted in more sensitive country-of-origin bias in Japan's beef imports. Kabe and Kanazawa (2012) employed a Markov-switching almost ideal demand system to identify the timing of structural change in Japanese meat market, and Soon and Thompson (2020) developed a time-varying Armington model with a Fourier-series approximation to explore how the U.S. BSE outbreaks affected Japanese beef demand.

While Japanese beef imports from the United States underwent a sudden drop followed by a gradual restoration of import purchases in the aftermath of BSE, the extent to which Japan's beef demand is characterized by a one-time shock—or, possibly, a nonlinear, more gradual shift—remains an important empirical question. This article contributes to the literature by developing a source-differentiated almost ideal demand system (SDAIDS) that embeds endogenous smooth transition functions to examine the time-varying nature of structural change caused by BSE in the United States, using monthly Japanese fresh/chilled and frozen beef import data from 1996 to 2016. The smooth transition function approach is appealing for several reasons. First, it provides a data-driven process to identify regime changes and transition starting and ending points (which are usually unknown), as opposed to previous studies, which needed to specify these points manually. Second, it allows a multiple regime switching and/or a nested linear specification to ensure flexible model fit and tests for or against various specifications including linear versus nonlinear, monotonic versus nonmonotonic, and two-regime versus multiple-regime transitions. Third, from a policy perspective, this flexibility is important because the reopening of the Japanese border to U.S. beef following BSE occurred gradually over a long period of time as restrictions were relaxed (e.g., beef derived from cattle aged 30 months or less).

Results show that a nonlinear structural change has occurred in the Japanese beef import market in the wake of BSE. Changing parameter estimates in the smooth transitional SDAIDS model result in changes in estimated beef import demand elasticities, which are important instruments for assessing trade policy impacts. Thus, the pre- and post-BSE elasticities shed light on the changing sensitivity of Japan's beef import purchases, the persistence of these changes, as well as the relative magnitudes of post-BSE elasticities compared to those in the pre-BSE period.

Empirical Model

The source-differentiated almost ideal demand system has been widely used to approximate import consumption behavior of goods from different origins (Yang and Koo, 1994; Andayani and Tilley, 1997; Henneberry and Hwang, 2007). In contrast to the Armington model, which suffers from restrictive assumptions of homotheticity and constant elasticity of substitution (Alston et al., 1990), the SDAIDS model ensures a flexible estimation of elasticity of substitution between goods from different origins. This is particularly important for international trade: Importing countries like Japan may perceive U.S. beef differently from Australian beef because of preferences and quality differences (Miljkovic and Jin, 2006; Obara, McConnell, and Dyck, 2010). Moreover, different transaction costs involved in international trade generate heterogeneous movements of import prices, making constant relative prices invalid in our case.

Following Yang and Koo (1994), the SDAIDS model defines the import share of good i imported from origin h at time t, w_{iht}, as a function of prices and expenditures:

$$w_{ih,t} = \alpha_{ih} + \sum_j \sum_k \gamma_{ihjk} \ln p_{jk,t} + \beta_{ih} \ln\left(\frac{E_t}{P_t}\right), \tag{1}$$

where $\ln P_t$ is the deflator price index,

$$\ln P_t = \alpha_0 + \sum_i \sum_h \alpha_{ih} \ln p_{ih,t} + \frac{1}{2} \sum_i \sum_h \sum_j \sum_k \gamma_{ihjk} \ln p_{ih,t} \ln p_{jk,t}, \tag{2}$$

where the subscripts i and j indicate goods $(i, j = 1, ..., M)$, subscripts h and k indicate origins of imported goods $(h, k = 1, ..., N)$, p_{ih} is the price of good i imported from origin h, and E is total expenditures on beef imports from all sources.[2]

Empirically, the SDAIDS model can be rewritten as

$$\begin{aligned} w_{ih,t} &= \alpha_{ih} + \sum_j \sum_k \gamma_{ihjk} \ln p_{jk,t} + \beta_{ih} (\ln E_t - \ln P_t) + \xi_{ih,t} \\ &= f(\boldsymbol{x}_t; \boldsymbol{\theta}) + \xi_{ih,t} \end{aligned} \tag{3}$$

where $\boldsymbol{x}_t$ contain all the explanatory variables; $\boldsymbol{\theta}$ is a vector of parameters to be estimated, including α, β, and γ; and $\xi_{ih,t}$ are the joint-normally distributed error terms with mean 0. The theoretical constraints imposed and tested in the SDAIDS model are

$$\begin{aligned} \text{adding-up:} \quad & \sum_i \sum_h \alpha_{ih} = 1, \\ & \sum_i \sum_h \gamma_{ihjk} = 0, \\ & \sum_i \sum_h \beta_{ih} = 0; \\ \text{homogeneity:} \quad & \sum_j \sum_k \gamma_{ihjk} = 0; \\ \text{symmetry:} \quad & \gamma_{ihjk} = \gamma_{jkih}. \end{aligned} \tag{4}$$

Due to the imposition of adding-up restriction in the demand system, the contemporaneous covariance matrix is singular. Hence, the last equation in the demand system is dropped for estimation purposes and the parameter estimates from the omitted equation are recovered using the theoretical restrictions. The model estimates are invariant to the decision of the dropped equation.

One concern about the SDAIDS model is that expenditures on a given partition of the utility function are likely to be endogenous when limiting demand estimation to a small subset of related goods (e.g., beef products). Ignoring this endogeneity can render the estimates biased and inconsistent. To address this issue, we specify a reduced-form equation of Japanese total beef import expenditures on a set of explanatory controls:

$$\begin{aligned} \ln E_t = c + c_1 BSE_t + \kappa \ln GDP_t + \textstyle\sum_i \sum_h \psi_{ih} \ln p_{ih,t} \\ + \psi_d \ln p_{d,t} + \psi_c \ln p_{c,t} + \psi_r \ln p_{r,t} + \zeta_t, \end{aligned} \tag{5}$$

where BSE_t is an indicator equal to 1 if the observation is in the period in which beef imports from the United States were banned due to BSE (January 2004–August 2006) and 0 otherwise; GDP_t is the gross domestic product (GDP) in Japan at time t; $p_{d,t}$, $p_{c,t}$, and $p_{r,t}$ denote the Japanese beef retail

[2] The SDAIDS model often suffers from degrees-of-freedom issues in empirical applications, depending on the number of imported goods and the number of origins per good. To ease this limitation, Yang and Koo (1994) proposed a restricted SDAIDS model by introducing the assumption of block substitutability, $\gamma_{ihjk} = \gamma_{ihj}, \forall\, k \in j \neq i$, such that cross-price effects of good j from either origin on the demand for good i from origin h are the same for all goods j regardless of their origins. We test and reject the assumption of block substitutability in Japanese beef import demand in favor of the unrestricted SDAIDS model.

price for domestic beef products, the Japanese consumer price index (CPI) of all commodities, and the Japanese real effective exchange rate at time t, respectively; and ζ_t denotes a well-behaved error term.

To test for endogeneity of beef import expenditures, we first estimate the reduced-form expenditure equation and retain the residuals $\hat{\zeta}_t$. We then add $\hat{\zeta}_t$ to each demand equation as an auxiliary variable and test the joint significance of the residual coefficients in the demand system. In cases where endogeneity is found to be an issue, we estimate jointly the demand system and expenditure equation using nonlinear SUR (seemingly unrelated regression)–iterative FGNLS (feasible generalized nonlinear least squares), which is equivalent to nonlinear full information maximum likelihood estimation, assuming the correct model specification with multivariate normal disturbances (Dhar, Chavas, and Gould, 2003).[3]

To add a nonlinear, nonmonotonic structure to the Japanese beef import demand system, we follow van Dijk, Teräsvirta, and Franses (2002) and Holt and Balagtas (2009) by incorporating a time-varying smooth transition function:

$$w_{ih,t} = f(\boldsymbol{x}_t;\theta_1) \times [1 - G(s_t;\lambda,c)] + f(\boldsymbol{x}_t;\theta_2) \times G(s_t;\lambda,c) + \xi_{ih,t}, \tag{6}$$

where $s_t = t/T$ is a time-transition variable. The transition function $G(\cdot)$ is a smooth, continuous function bounded between 0 and 1 according to s_t. The SDAIDS model in equation (6) can be thought of as a two-regime switching model. Each of the two regimes associated with the extreme value of $G(\cdot) = 0$ and $G(\cdot) = 1$, where the transition from one regime to another is continuous and smooth. θ_1 and θ_2 are the parameter sets identifying the two regimes of the SDAIDS model, λ is the speed-of-adjustment parameter that determines how quickly the model shifts regimes, and c is the threshold parameter that defines the point at which the transition is 50% complete or symmetric. Consumer preferences represented by the underlying model parameters are permitted to vary over time as the transition function changes.

We begin with a two-regime model and find it restrictive since it essentially forces demand in the middle regime to be a weighted combination of demand in the pre- and post-BSE regimes. We argue that a three-regime framework is more appropriate. Particularly, the three distinct policy regimes are identified as the pre-BSE period (up to December 2003), the BSE-ban period (January 2004–August 2006), and the post-BSE recovery period (post September 2006). Our decision is made based on two points. First, the observed data support three regimes over 1996–2016, which is aligned with the usual policy setting (pre-treatment, treatment, and post-treatment periods). Second, given that the smooth transitional SDAIDS model is semiparametric, we face a tradeoff between model complexity and effective degrees of freedom. As the model becomes more complex, it will be more prone to overfitting.

As a result, a three-regime framework is employed by adding a second nonlinear component, yielding:

$$\begin{aligned} w_{ih,t} = {} & f(\boldsymbol{x}_t;\theta_1) \times [1 - G_1(s_t;\lambda_1,c_1)] \\ & + f(\boldsymbol{x}_t;\theta_2) \times [G_1(s_t;\lambda_1,c_1) - G_2(s_t;\lambda_2,c_2)] \\ & + f(\boldsymbol{x}_t;\theta_3) \times G_2(s_t;\lambda_2,c_2) + \xi_{ih,t} \end{aligned} \tag{7}$$

If it is assumed that $c_1 < c_2$, the parameters in this model change smoothly from θ_1 via θ_2 to θ_3 for increasing values of s_t, as the first function G_1 changes from 0 to 1, followed by a similar change in G_2 (van Dijk, Teräsvirta, and Franses, 2002).

An important step in estimation is the selection of a transition function that better represents the observed situation. In this study, we consider two standard specifications (Lin and Teräsvirta, 1994;

[3] Another concern is the assumption of price exogeneity. Using a beef import demand model conditional on total beef import expenditures, we assume that Japanese consumers respond to the given relative import price changes. This assumption may not hold if the model extends to both domestic produced and imported beef commodities where domestic beef (wagyu and non-wagyu breeds) prices could be endogenous in response to varying production costs.

Teräsvirta, 1994; van Dijk, Teräsvirta, and Franses, 2002). The first is the logistic smooth transition (LSTR) function:

$$G(s_t,\lambda,c) = \left\{1 + \exp\left[-\lambda\left(\frac{s_t - c}{\sigma_{s_t}}\right)\right]\right\}^{-1}, \tag{8}$$

where the speed-of-adjustment parameter, λ, is expressed as $\lambda = \exp(-\lambda^*)$ to ensure that λ is positive by definition, and σ_{s_t} is the standard deviation of the normalized trend variable to ensure that λ is unit invariant. The threshold parameter, c, is the centrality parameter, indicating the point at which the transition is 50% complete. The LSTR function changes monotonically from 0 to 1 as s_t increases. As $\lambda \to 0$, $G(\cdot)$ in equation (8) is effectively linear in s_t, whereas when $\lambda \to \propto$, the LSTR function in equation (8) becomes a Heaviside indicator function that equals 0 if $s_t < c$, and 1 otherwise. Consequently, a change of $G(\cdot)$ from 0 to 1 is instantaneous at $s_t = c$ when $\lambda \to \propto$.

An alternative is the exponential smooth transition (ESTR) function:

$$G(s_t,\lambda,c) = 1 - \exp\left[-\lambda\left(\frac{s_t - c}{\sigma_{s_t}}\right)^2\right], \tag{9}$$

where the parameters λ and σ_{s_t} are defined as in equation (8) except that the threshold parameter, c, indicates the point at which the transition is symmetric. Structural change implied by the ESTR function in equation (9) is nonmonotonic and symmetric around c. For $\lambda \to 0$ and $\lambda \to \propto$, $G(\cdot)$ approaches 0 and 1, respectively. When $s_t \to \pm\propto$, $G(\cdot) \to 1$, whereas $s_t = c$, $G(\cdot) = 0$.

Using the three-regime SDAIDS model, the estimated expenditure (η_{ih}) and uncompensated price elasticities (ε^u_{ihjk}) are obtained as follows:

$$\eta_{ih} = \frac{\kappa}{\bar{w}_{ih}}\left\{\beta^1_{ih} \times (1 - G_1) + \beta^2_{ih} \times (G_1 - G_2) + \beta^3_{ih} \times G_2\right\} + 1; \tag{10}$$

$$\begin{aligned}\varepsilon^u_{ihjk} = \frac{1}{\bar{w}_{ih}}&\left\{\left[\gamma^1_{ihjk} + \beta^1_{ih}\left(\psi_{jk} - \alpha^1_{jk} - \sum_r\sum_m \gamma^1_{jkrm}\ln\bar{p}_{rm}\right) \times (1 - G_1)\right]\right.\\ &+ \left[\gamma^2_{ihjk} + \beta^2_{ih}\left(\psi_{jk} - \alpha^2_{jk} - \sum_r\sum_m \gamma^2_{jkrm}\ln\bar{p}_{rm}\right) \times (G_1 - G_2)\right]\\ &\left.+ \left[\gamma^3_{ihjk} + \beta^3_{ih}\left(\psi_{jk} - \alpha^3_{jk} - \sum_r\sum_m \gamma^3_{jkrm}\ln\bar{p}_{rm}\right) \times G_2\right]\right\} - \delta_{ihjk};\end{aligned} \tag{11}$$

where $\bar{w}$ and $\bar{p}$ are the sample mean budget shares and prices; κ and ψ_{jk} are the expenditure and price coefficients from the auxiliary expenditure equation; δ_{ihjk} is the Kronecker delta, which equals 1 if $i = j$, $h = k$, and 0 otherwise.

Data

Japan's monthly beef import data from January 1996 to December 2016 are collected from the Global Trade Atlas Database.[4] Four exporting countries/regions are included in the smooth transitional SDAIDS model. Three major source countries are selected based on the significance of their market shares in Japanese beef imports: the United States (USA), Australia (AUS) and New Zealand (NZL). The remaining exporting countries are combined into a composite rest of the world (ROW) region. Beef products are distinguished by product type: fresh/chilled beef (HS 0201) and frozen beef (HS 0202), which account for about 85% of Japan's total beef and offal imports during

[4] Data are retrieved from https://www.gtis.com/gta/, which requires subscription.

the sample period. Beef offal imports are excluded in the model estimation, due to the problem of representing heterogeneous offal products using a single price.[5] Table 1 reports summary statistics of model data.

Import prices for beef products from different origins are not publicly reported. Hence, unit values are employed as a proxy for import prices, obtained by dividing import value by import quantity (Yang and Koo, 1994; Andayani and Tilley, 1997; Henneberry and Hwang, 2007; Grant, Lambert, and Foster, 2010). To minimize aggregation bias of unit values, we compute the Stone price index for each source-differentiated beef product at the HS four-digit level using import values and quantities from the HS six-digit products, with the weights being average import shares in each HS four-digit product sector and time period.[6] When Japan's beef import prices from the United States are missing during the BSE-ban period due to 0 trade, we approximate import prices using monthly maximum plus 1 standard deviation.[7]

We impose weak separability between imported and domestic beef products in the Japanese market in this study for two reasons. First, imported beef has distinct attributes compared to domestic beef (e.g., wagyu beef). Further, imported and domestic beef products are marketed and consumed via different channels (Peterson and Chen, 2005; Obara, McConnell, and Dyck, 2010). Second, we encountered significant data limitations associated with obtaining the quantity of monthly domestic beef purchases in Japan. One solution we considered was to use more aggregated data to match available domestic beef expenditures and quantities consumed in Japan. However, estimation of the smooth transitional SDAIDS model was not feasible due to insufficient degrees of freedom.

In the reduced-form expenditure regression (equation 5), quarterly GDP (constant 2010 $U.S. dollars, seasonally adjusted),[8] monthly CPI (2010 base year, seasonally adjusted) and monthly real effective exchange rates (in 2010 base year) are collected from the World Bank Global Economic Monitor.[9] Monthly domestic beef retail prices (in 2010 base year) are collected from Japanese Agriculture and Livestock Industries Corporation (ALIC).[10]

Empirical Results

Tests for Model Assumptions

The demand system contains prices of beef imported from different origins in each equation. To determine whether the SDAIDS model is valid in estimating Japanese beef import demand compared to the canonical AIDS model, we test the assumption of product aggregation and block separability. Results show that the null hypothesis of aggregated beef imports with no differentiation of exporting sources is rejected at the 1% significance level. Next, the assumption of block substitutability is tested and rejected at the 1% significance level, suggesting that there are important cross-price

[5] Before December 2003, fresh/chilled beef cuts accounted for 42% of Japan's total beef and offal imports, frozen beef cuts for 44%, and other edible offal meat for 14%. After the U.S. BSE outbreak, Japan began importing more frozen beef (52%), less fresh/chilled beef (38%), and even less edible offal meat (10%).

[6] For fresh/chilled beef imports, we include the Harmonized System (HS) codes 020110, 020120 and 020130. For frozen beef imports, we include HS codes 020210, 020220 and 020230.

[7] The traditional approach is to approximate missing prices by sample/cluster mean unit values. Alternatively, Muhammad (2013) proposes an iterative procedure to estimate choke prices. We approximated the missing prices using both monthly mean and maximum unit values and found no significant differences in our main conclusions. We thus decided to use a monthly maximum plus 1 standard deviation to reflect an approximation of the prohibitive effect of the BSE outbreaks on U.S. beef import prices and seasonal variability.

[8] We apply the cubic spline interpolation to get the estimates of monthly GDP data from quarterly GDP data for the empirical analysis.

[9] https://datacatalog.worldbank.org/dataset/global-economic-monitor

[10] http://www.alic.go.jp/english/

Table 1. Summary Statistics of Model Variables

	Import Expenditures ($ millions)			Import Shares (%)			Prices ($/kg)		
	Pre-BSE	**BSE Ban**	**Post-BSE**	**Pre-BSE**	**BSE Ban**	**Post-BSE**	**Pre-BSE**	**BSE Ban**	**Post-BSE**
Chilled									
USA	66.22	0.00	34.93	34.10	0.00	16.10	6.30	9.20	6.98
AUS	56.52	99.51	76.50	29.60	62.42	38.40	3.81	5.66	6.43
NZL	1.88	2.54	3.84	0.98	1.62	1.90	4.99	6.80	7.34
ROW	1.71	1.32	2.48	0.90	0.82	1.21	6.72	7.83	7.39
ALL	126.30	103.37	117.75	65.64	64.86	57.70	4.85	5.70	6.60
Frozen									
USA	44.02	0.00	22.34	21.80	0.00	10.10	2.93	6.43	4.41
AUS	18.70	44.16	52.72	9.74	27.64	25.60	1.88	2.78	3.49
NZL	2.78	9.61	7.19	1.40	5.91	3.54	2.61	3.38	4.32
ROW	2.62	2.12	6.52	1.37	1.30	3.04	3.40	5.07	4.99
ALL	68.13	55.88	88.77	34.36	34.85	42.31	2.50	2.92	3.71

Notes: The pre-BSE period is January 1996–December 2003; the BSE-ban period is January 2004–August 2006; and the post-BSE period is September 2006–December 2016. Import expenditures are monthly average expenditures on the corresponding beef imports. Import shares are sample average expenditure shares of the corresponding beef imports. Prices are sample average unit values per kilogram for the corresponding beef imports.

Table 2. Model Fit, Test Statistics, and Transition Function Parameter Estimates for the Japanese Beef Import Demand Models

	Basic SDAIDS	Two-Regime SDAIDS	Two-Regime SDAIDS	Three-Regime SDAIDS
	1	2	3	4
No. of parameters	49	86	86	123
Log likehood	4,939	5,254	5,258	5,444
System AIC	−9,780	−10,336	−10,343	−10,641
System BIC	−9,607	−10,033	−10,040	−10,207
LR test statistic	–	638.38	645.41	1,017.58
		[0.001]	[0.001]	[0.001]
Smooth transition function	–	logistic	exponential	exponential & logistic
λ_1^*	–	−1.318	−7.966	−1.117
		(0.267)	(0.259)	(0.199)
λ_2^*	–	–	–	1.131
				(0.111)
c_1	–	4.855	−1.157	0.077
		(0.641)	(0.394)	(0.015)
c_2	–	–	–	0.680
				(0.013)
Expenditure regression	YES	YES	YES	YES

Notes: Number of observations in each model is 2,016. In column 4, the exponential transition function has a speed-of-adjustment parameter $\lambda_1 = \exp(-1.117) = 0.327$ and a symmetry parameter, $c_1 = 0.077$; the logistical transition function has a speed-of-adjustment parameter $\lambda_2 = \exp(1.131) = 3.099$ and a centrality parameter, $c_2 = 0.68$. AIC denotes the Akaike information criterion. BIC denotes the Bayesian information criterion. Smaller values (i.e., more negative) of the AIC and BIC indicate a better model fit. LR denotes the likelihood ratio test. Test statistic p-values are in brackets. Standard errors are in parentheses.

substitution effects between the source-differentiated fresh/chilled and frozen beef imports in the Japanese market. Thus, the data support an unrestricted SDAIDS model.[11]

Estimates from the demand model without accounting for the potential endogeneity of import expenditures may be biased and inconsistent. To address this issue, we include the residual estimates from the reduced-form expenditure equation in the SDAIDS model and apply the Wu–Hausman statistic to test for significance of the residual coefficients (Wu, 1973; Hausman, 1978). The null hypothesis of strict exogeneity of the expenditure term (insignificant correlation of the expenditure variable with the demand system error terms) is rejected at the 1% significance level.[12] As a result, we estimate the demand system equations and the expenditure equation jointly, leading to 49, 86, and 123 parameters to be estimated for the baseline, two-regime, and three-regime switching SDAIDS models, respectively.

Table 2 reports key measures of model fit, test statistics, and transitional function parameter estimates for the Japanese beef import demand system under various specifications. The likelihood ratio (LR) tests for the null hypothesis of no structural change (i.e., $\theta_2 = 0$ in equation 6 or $\theta_2 = \theta_3 = 0$ in equation 8) are rejected at the 1% significance level, suggesting that parameter estimates in the Japanese beef import demand model vary significantly over the sample period.

[11] After imposing homogeneity and symmetry restrictions in the SDAIDS model, the test statistic for product aggregation in the system is $\chi^2(36) = 126288$; the test statistic for block separability is $\chi^2(24) = 234$; the test statistic for block substitutability is $\chi^2(12) = 91$.

[12] After imposing homogeneity and symmetry restrictions, the test statistic for expenditure exogeneity is $\chi^2(7) = 74$ in the SDAIDS model.

Further, the LR tests for the null hypothesis of a two-regime switching demand model are rejected at the 1% significance level in favor of a three-regime smooth transition SDAIDS model.

By comparing models in terms of the log-likelihood value, system Akaike information criterion (AIC), and Bayesian information criterion (BIC), we conclude that the three-regime SDAIDS model provides a better model fit and ability to capture both the instantaneous and longer-run structural change caused by the U.S. BSE event. Thus, in the remainder of the paper, we focus on the three-regime switching SDAIDS model. The estimated coefficients of the three-regime SDAIDS model and the smooth transition functions are provided in Table S1 and Figures S1 and S2 in the online supplement (see www.jareonline.org).[13]

Results show that over half of α_{ih} and several β_{ih} changed drastically during the BSE-ban period and did not fully recover to the pre-BSE levels even after the ban on U.S. beef exports to Japan was lifted.[14] The joint test statistics for $\alpha_{ih}^1 = \alpha_{ih}^3$, $\beta_{ih}^1 = \beta_{ih}^3$, and $\gamma_{ihjk}^1 = \gamma_{ihjk}^3$ (pre- versus post-BSE estimates) indicate that the null hypothesis of no parameter changes of α_{ih}, β_{ih} and γ_{ihjk} are all significantly rejected. Thus, results suggest the presence of a persistent, nonlinear structural change in Japanese demand for source-differentiated beef imports in the wake of BSE.

Elasticities in SDAIDS Model

Expenditure Elasticities

To examine how changes in preference parameters affected Japan's expenditure elasticities of beef demand (η_{ih} in equation 10), Figure 2 presents time-series plots of estimates for Japanese beef imports from the United States, Australia, and New Zealand using sample average budget shares and prices. The panels in the lefthand column display the estimated expenditure elasticities of Japanese demand for chilled beef imports from selected exporting countries, and those in the righthand column depict the corresponding estimated expenditure elasticities of Japanese demand for frozen beef imports. As a reference, we report the 5% lower bound and 95% upper bound confidence intervals in the grey-shaded area in each panel.

With the exception of frozen beef imports from New Zealand, all expenditure elasticities are positive and statistically significant at the 5% level, suggesting that Japanese consumers treat source-differentiated beef imports as normal goods (Figure 2). Several trends are worth highlighting. First, prior to the BSE outbreaks in late 2003, Japanese expenditure elasticities were relatively stable for beef imports from the United States, while that for Australian and New Zealand beef imports showcased more variation. However, this pattern reversed early in 2004—Japanese expenditure elasticities for U.S. beef imports became more volatile compared to those for Australian and New Zealand beef imports. Second, pre-BSE expenditure elasticities for chilled beef imports from the United States were relatively less elastic with respect to changes in Japanese expenditures compared to those from non-U.S. suppliers. In contrast, expenditure elasticities for frozen beef imports from the United States were much more expenditure elastic. It seems that Japanese consumers considered chilled beef from New Zealand and frozen beef from the United States to be more luxurious than those from other sources.

Third, expenditure elasticities gradually declined as budget shares on U.S. beef imports recover from the 2003 BSE events. None of them, however, reached their pre-BSE levels. Japanese expenditure elasticities for U.S. chilled (frozen) beef increased (dropped) significantly, by 62.5% (33.2%), while those for Australian chilled and frozen beef barely changed. The differences in the pre-versus-post comparison of expenditure elasticities for Japan's beef imports from the United

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[13] Because we incorporated two different transition functions into the SDAIDS model, it is not straightforward to see the transition procedure when viewing the transition function separately. One needs to combine each transition function with the corresponding demand parameters in order to fully understand how import demand has shifted. Therefore, it is suggested that information in Figures A1 and A2 should be jointly considered.

[14] For example, $\alpha_{\text{USA, chilled}}$, $\alpha_{\text{USA, frozen}}$, $\alpha_{\text{AUS, chilled}}$, $\alpha_{\text{AUS, frozen}}$; $\beta_{\text{USA, chilled}}$, $\beta_{\text{USA, frozen}}$, $\beta_{\text{AUS, chilled}}$, $\beta_{\text{AUS, frozen}}$.

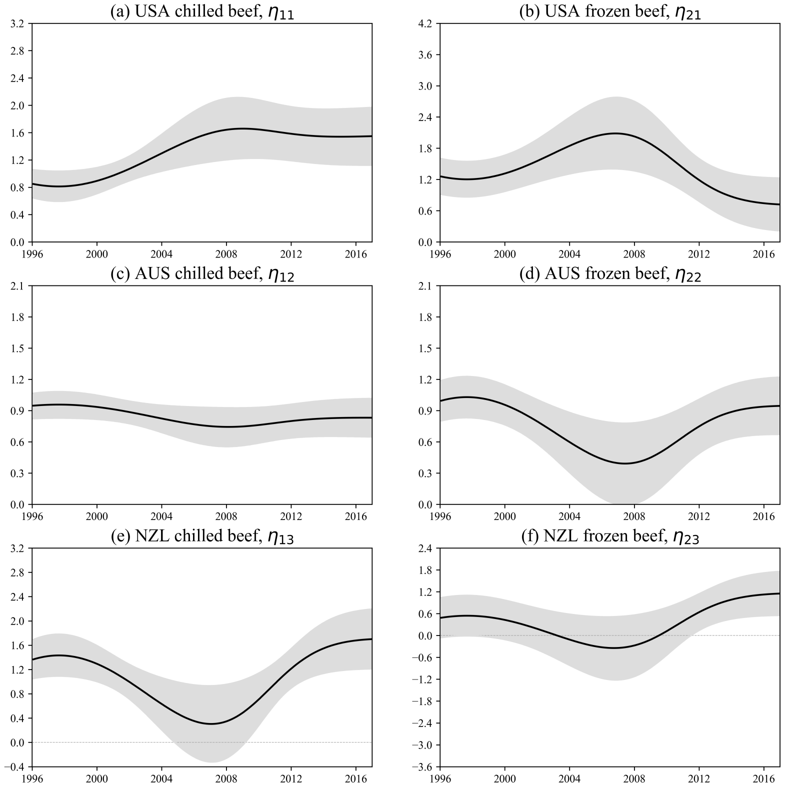

Figure 2. Time-Series Plots of Expenditure Elasticities for Japanese Beef Imports

Notes: Time-series estimates and 95% confidence intervals (grey areas) are derived from the three-regime smooth transition SDAIDS model for Japanese beef imports.

States are statistically significant at the 5% level. For any given increase in expenditures, Japan has become more willing to increase its demand for chilled beef rather than frozen beef from the United States.

Own-Price Elasticities

Figure 3 presents the time-series plots of uncompensated own-price elasticities (ε^{u}_{ihih} in equation 11) for Japanese beef imports from the United States, Australia, and New Zealand, with the panels in the lefthand column for source-differentiated chilled beef imports and those in the righthand column for source-differentiated frozen beef imports. With the exception of New Zealand beef demand in the BSE ban period, all own-price elasticities were negative and statistically significant at the 5% level. Japanese own-price elasticities for U.S. beef imports were quite stable in the pre-BSE period but more volatile in the post-BSE era, whereas own-price elasticities for Australian and New Zealand beef imports varied more were varied more before the outbreaks but were more stable afterwards. Results show that own-price elasticities for chilled beef imports from all three major origins are much more price elastic while that for frozen beef imports are less sensitive to own-price changes.

Changes in parameter estimates (and budget shares) have led to persistent changes in Japanese own-price beef import demand elasticities. Therefore, it had yet been able to reach the corresponding pre-BSE levels at the end of the sample period. Japan's own-price elasticities for U.S. (Australian) chilled beef imports declined dramatically, by 55.5% (31.7%) in absolute terms, whereas own-price elasticities for U.S. frozen beef imports nearly tripled in absolute terms. The pre-versus-post changes in own-price elasticities are statistically significant at the 5% level, implying that, after 4–6 years of

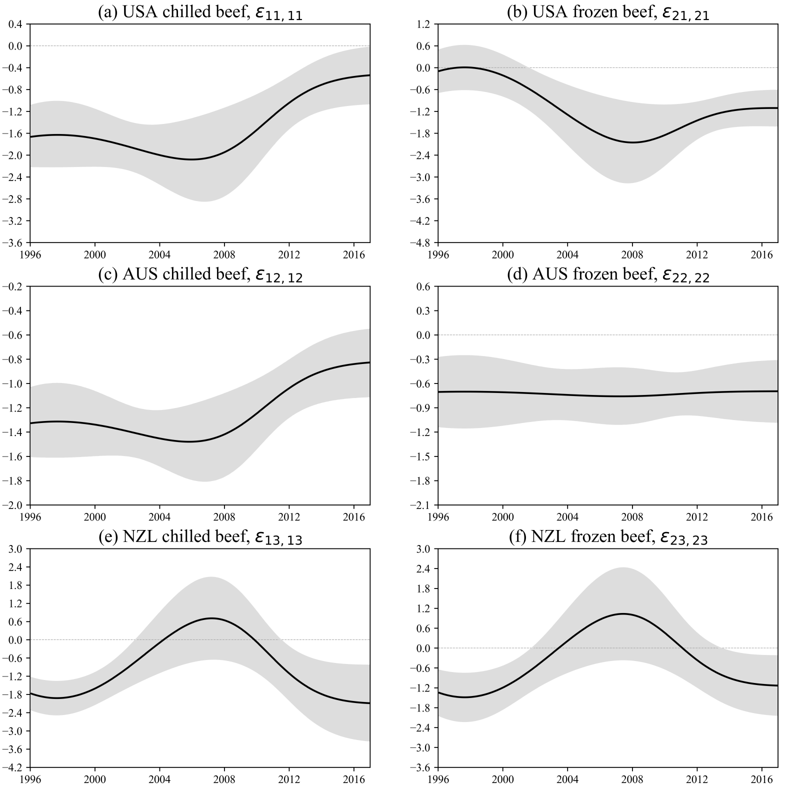

Figure 3. Time-Series Plots of Own-Price Elasticities for Japanese Beef Imports

Notes: Time-series estimates and 95% confidence intervals (grey areas) derived from the three-regime smooth transition SDAIDS model for Japanese beef imports.

recovery following the BSE outbreak, Japanese demand for U.S. (Australian) beef products became much less (more) responsive to own-price changes.

Cross-Price Elasticities

To examine the competitiveness of major beef exporters in Japan, we turn our attention to cross-price elasticities (ε^u_{ihjk} in equation 11). Most of the elasticity estimates are statistically significant at the 5% level, suggesting the presence of cross-product and group substitution effects in the Japanese market for imported beef. Different signs in the estimates indicate a variety of substitute and complementary relationships among different sources. Over half of cross-price elasticities have undergone noticeable changes in the post-BSE period, indicating alternative sourcing of beef imports among competing suppliers in the Japanese market. Below, we first discuss cross-price elasticities for different beef products from the same origin (across product categories for a given origin country), then cross-price elasticities for beef products from different origins (across origins and product categories).

Figure 4 presents time-series plots of within-origin cross-price elasticities (ε^u_{ihik}). Interestingly, before late 2003, within-origin cross-price elasticities were all negative, meaning that different beef products imported from the same origin tended to have a complementary relationship in the Japanese market. In other words, any price decrease of chilled beef imported from a given source country was associated with an increase in the demand for frozen beef imported from the same source. However, after the BSE crisis, some estimates become positive, showing a competing relationship between different beef products imported from the same origin in the Japanese market.

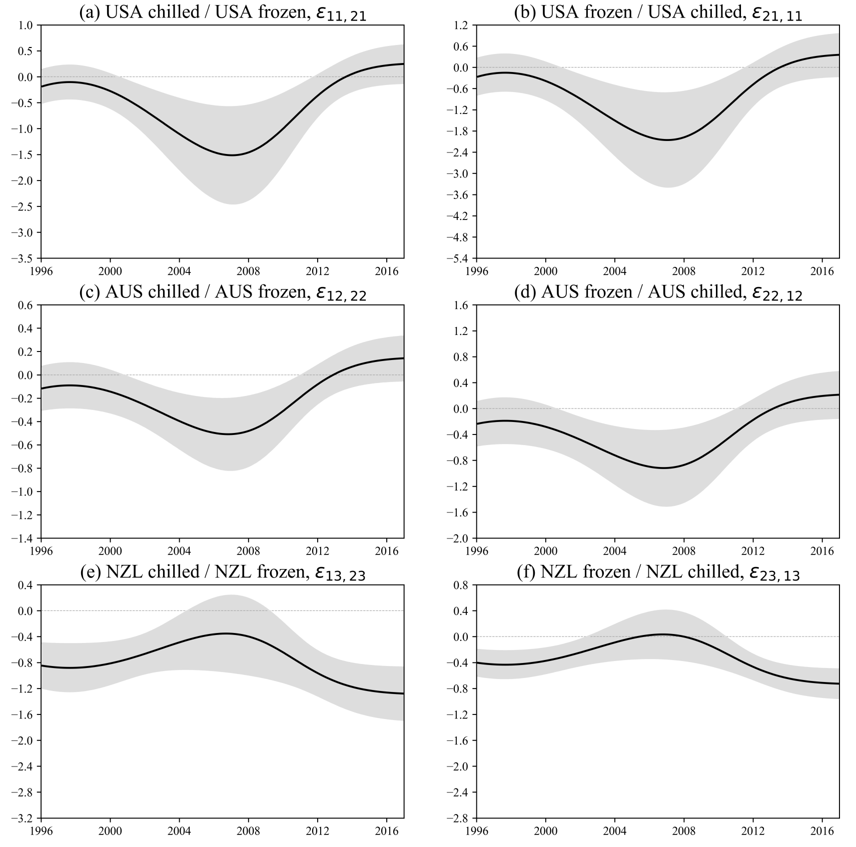

Figure 4. Time-Series Plots of Within-Origin Cross-Price Elasticities for Japanese Beef Imports

Notes: Time-series estimates and 95% confidence intervals (grey areas) derived from the three-regime smooth transition SDAIDS model for Japanese beef imports.

To economize on space, Figure 5 presents time-series plots of across-origin cross-price elasticities (ε^{u}_{ihjk}) involving only the United States and Australia. Before the BSE outbreak, Japanese cross-price elasticities for U.S. beef imports were relatively stable with respect to price changes of Australian beef, while those for Australian beef imports in response to the change in U.S. beef prices show larger variation. However, a reversal occurred in the aftermath of BSE. By comparing each subfigure in Figure 5, we find that Japanese demand for U.S. beef was generally more responsive to price increases of Australian beef than Japanese demand for Australian beef in response to price changes of U.S. beef. In addition, before the BSE outbreak, most cross-price elasticities for U.S. and Australian beef imports (except for the frozen beef, Figure 5(d) and (h)) were positive. The substitution effects even intensified in 2004–2007, likely owing to the prior ban on U.S. beef imports. For example, Japanese demand for U.S. beef imports was more than twice as sensitive in response to price changes of Australian beef in the ban period, while demand for Australian beef was slightly more responsive to price changes of U.S. beef. After 4–6 years of market recovery, cross-price elasticities have now become less positive and even negative in most cases (though statistically insignificant from 0), implying that Japanese beef imports from the United States and Australia have become more complementary over time.

To put the analysis in context, one reviewer noted that it would be interesting to compare our post-BSE elasticity estimates (using data from 2011–2016) to similar studies (see, e.g., Muhammad, Countryman, and Heerman, 2018; Soon and Thompson, 2020). A few differences in elasticity estimates arise due to different datasets, sample periods, and frequencies and methodologies applied.

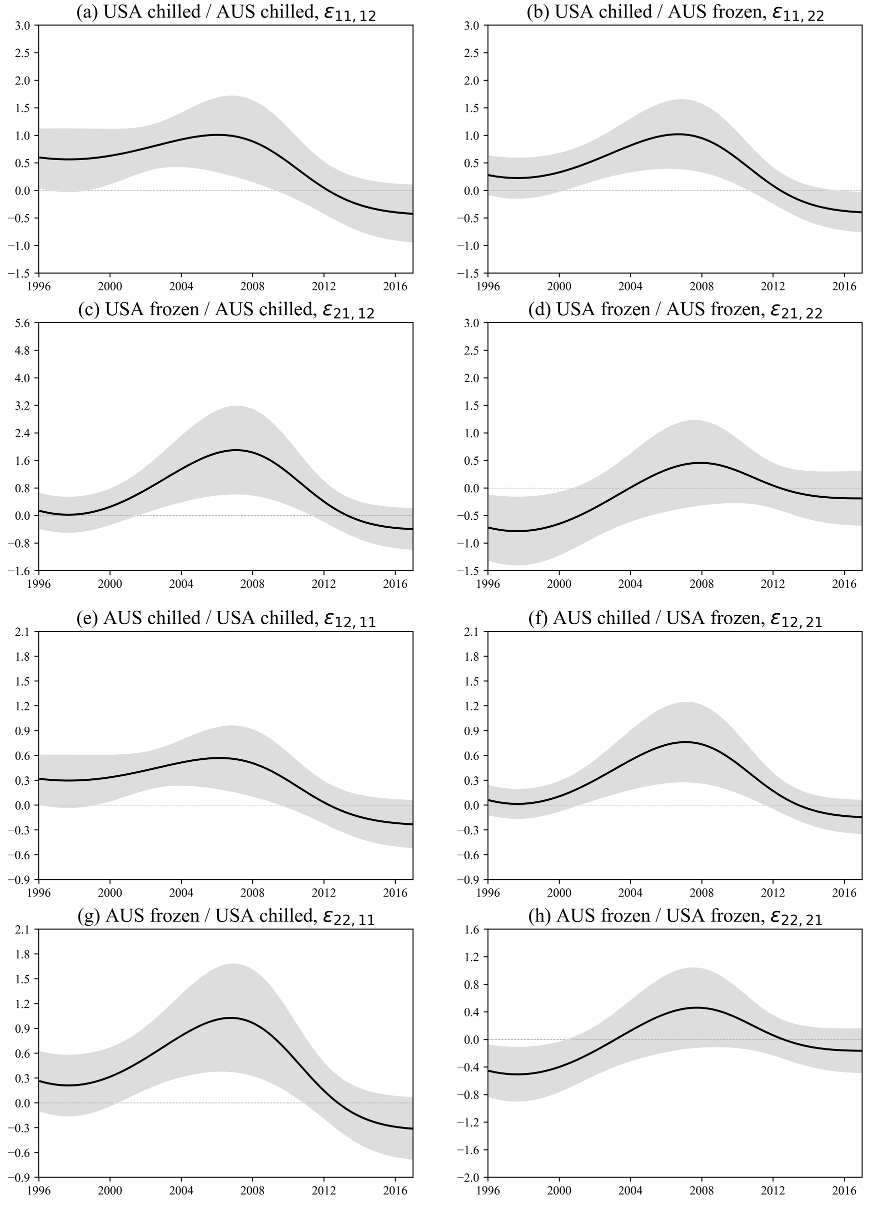

Figure 5. Time-Series Plots of (Selected) Cross-Price Elasticities for Japanese Beef Imports

Notes: Time-series estimates and 95% confidence intervals (grey areas) derived from the three-regime smooth transition SDAIDS model for Japanese beef imports.

We find that expenditure elasticities for U.S. beef (1.55 in chilled, 0.94 in frozen) are much larger than those for Australian beef (0.78 in chilled, 0.88 in frozen), reflecting the increasing recovery of U.S. beef imports (especially for chilled beef) during the sample period. Own-price elasticities for U.S. beef (-0.82 in chilled, -1.27 in frozen) and Australian beef (-0.93 in chilled, -0.77 in frozen) are relatively more elastic than the estimates from Muhammad, Countryman, and Heerman (2018). Additionally, Japanese demand for beef imports is especially responsive to price changes of chilled versus frozen beef—an interesting feature that has not been addressed in previous studies. Together, we find that the U.S. BSE outbreaks not only stimulated Japan's demand for each beef variety from different origins but also the combination of different beef varieties from the same origin.

Concluding Remarks

This study developed an SDAIDS model augmented with smooth transition functions to test for a possibly nonlinear structural shock caused by the U.S. BSE outbreak and to evaluate its impact on Japanese beef imports using estimated demand elasticities. Although assessing the consequences of animal disease outbreaks is challenging, this paper proposed a novel method for empirical implementation of import demand estimation. Several important findings are summarized below. First, the BSE incident led to a persistent and long-lasting impact on source-differentiated beef import demand in Japan. While restrictions on U.S. beef exports were lifted after August 2006, it took much longer (4–6 years) for Japanese consumers to rebuild confidence in U.S. beef products. Second, the BSE-induced structural change appears to be nonlinear. Hence, import demand estimates without consideration of a nonlinear structural break could result in incomplete or misleading policy implications. Third, economically important changes in the magnitude and variability of estimated elasticities emerged in the wake of BSE. Import demand substitutability varied a lot over time owing to parameter changes. In particular, during the recovery period, Japan has become more likely to purchase fresh/chilled beef but less likely to purchase frozen beef from the United States as expenditures increase. Given that consumers generally prefer fresh/chilled beef over frozen beef, it seems reasonable that they tend to be more expenditure-sensitive to the former. In contrast, the model estimated very little change in Japanese expenditure elasticities for Australian beef imports. Fourth, Japanese demand for U.S. (and Australian) chilled beef imports has become less own-price elastic, compared to the demand for U.S. frozen beef imports, which is more own-price elastic.

Our results show that because of the persistent and long-term consequences of BSE, Japanese beef import demand entered a new regime around 2010 that has not fully recovered to the pre-BSE situation. Therefore, future studies characterizing this market should focus on the post-BSE estimates for policy evaluation. For example, one could use our post-BSE import demand elasticities to evaluate the competitiveness of U.S. beef exports in the Japanese market due to the establishment of Japanese free trade agreements (FTAs) with other major exporting countries (Muhammad, Countryman, and Heerman, 2018, see). One could also use them in general equilibrium models to assess the impact of bilateral and multilateral trade agreements on regional and global meat trade, particularly because CGE results are often sensitive to import demand elasticities (see, e.g., Hertel et al., 2007; Hillberry and Hummels, 2013).

In a broader context, our results have important implications for the meat protein trade and the recovery of meat imports following significant food safety events. For example, since the April 2009 H1N1 Influenza virus (swine flu) outbreaks were detected first in the United States and spread to other nations, some key pork markets—including China, Philippines, Russia, and Ukraine—imposed bans on pork imports from the United States. African swine fever and H5N1 Avian Influenza virus (bird flu) outbreaks have been met with similarly strict import requirements or bans by major pork and poultry importers, respectively. With increasing protein demand from middle-income and emerging countries, an important policy question is how import demand purchases and substitution patterns for beef, pork, and poultry products will recover in the wake of food safety outbreaks. This study provides some intriguing clues using the 2003 U.S. BSE event.

Finally, several caveats are worth noting. First, the SDAIDS model employed in this study does not include Japanese domestic beef purchases. Although Japan is not a major beef producer, it is possible that the BSE events may have caused Japanese consumers to substitute imported products with what they considered to be safer domestic products. Future studies should consider domestic expenditures where data are available. Second, the BSE outbreaks may have affected Japanese demand for other meat varieties (e.g., pork, poultry, and seafood products). Future studies could develop an unconditional demand system of a wider set of meat varieties from different origins and explore meat substitution. Third, future studies should consider incorporating a flexible number of transition regimes suited for the empirical demand system, provided there are sufficient degrees of freedom in estimation. Nonetheless, this study contributes to the import demand and structural change literature by examining the impact of the U.S. BSE outbreaks on Japanese beef import demand and provides important policy implications about the recovery of meat import demand and substitution patterns in the wake of food safety concerns.

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Online Supplement: Estimating Structural Change in the Japanese Beef Import Market in the Wake of BSE: A Smooth Transition Approach

Xin Ning, Jason H. Grant, Everett B. Peterson

Table S1. Estimated Coefficients from the Expenditure Regression for the Japanese Beef Import Demand Model

	Coefficient	**Std. Err.**
Constant	0.046***	(0.014)
ln(GDP)	2.622***	(0.677)
BSE ban	−0.453***	(0.053)
Japanese domestic beef retailer price	−0.908**	(0.431)
Japanese consumer price index	3.316**	(1.437)
Japanese real effective exchange rate	0.451***	(0.112)
ln(price of U.S. chilled beef)	0.595***	(0.131)
ln(price of Australian chilled beef)	0.095	(0.213)
ln(price of New Zealand chilled beef)	−0.277	(0.170)
ln(price of rest-of-world chilled beef)	0.078*	(0.045)
ln(price of U.S. frozen beef)	0.101	(0.085)
ln(price of Australian frozen beef)	−0.256*	(0.135)
ln(price of New Zealand frozen beef)	0.521***	(0.101)
ln(price of rest-of-world frozen beef)	−0.025	(0.024)

Table S2. Estimated Coefficients from the Demand System Regression for the Japanese Beef Import Demand Model

	USA Chilled Beef	AUS Chilled Beef	NZL Chilled Beef	ROW Chilled Beef	USA Frozen Beef	AUS Frozen Beef	NZL Frozen Beef	ROW Frozen Beef
α^1	0.433***	0.253***	0.011***	−0.007***	0.195***	0.077***	0.036***	0.001
	(0.026)	(0.026)	(0.002)	(0.003)	(0.022)	(0.020)	(0.008)	(0.004)
α^2	−0.414***	0.895***	0.008*	0.035***	−0.133*	0.511***	0.057***	0.040***
	(0.149)	(0.124)	(0.004)	(0.009)	(0.074)	(0.082)	(0.016)	(0.010)
α^3	0.318***	0.232***	0.027***	0.007***	0.145***	0.196***	0.034***	0.040***
	(0.033)	(0.031)	(0.002)	(0.003)	(0.026)	(0.024)	(0.008)	(0.006)
β^1	−0.015*	−0.006	0.002***	0.006***	0.011	0.002	−0.005*	0.006***
	(0.008)	(0.009)	(0.001)	(0.001)	(0.008)	(0.008)	(0.003)	(0.002)
β^2	0.122***	−0.076***	−0.015***	−0.003	0.136***	−0.119***	−0.035***	−0.009**
	(0.032)	(0.028)	(0.004)	(0.003)	(0.024)	(0.020)	(0.009)	(0.004)
β^3	0.046***	−0.025**	0.004***	−0.003***	−0.014	−0.005	0.002	−0.004
	(0.013)	(0.013)	(0.001)	(0.001)	(0.013)	(0.011)	(0.004)	(0.003)
$\gamma^1_{\text{USA chilled}}$	−0.130**	0.116*	−0.003	0.016**	−0.024	0.042	−0.022	0.004
	(0.063)	(0.060)	(0.006)	(0.007)	(0.035)	(0.037)	(0.018)	(0.006)
$\gamma^1_{\text{AUS chilled}}$		−0.122**	0.018***	−0.012*	0.007	−0.038	0.037**	−0.006
		(0.058)	(0.007)	(0.006)	(0.033)	(0.035)	(0.018)	(0.006)
$\gamma^1_{\text{NZL chilled}}$			−0.013***	−0.002*	0.012***	0.001	−0.015***	0.002***
			(0.004)	(0.001)	(0.004)	(0.005)	(0.003)	(0.000)
$\gamma^1_{\text{ROW chilled}}$				−0.002	0.005	−0.005	−0.000	0.000
				0.002	(0.004)	(0.005)	(0.003)	(0.001)
$\gamma^1_{\text{USA frozen}}$					0.134***	−0.100**	−0.029**	−0.005
					(0.041)	(0.039)	(0.014)	(0.006)
$\gamma^1_{\text{AUS frozen}}$						0.06	0.038**	0.001
						(0.043)	(0.016)	(0.006)
$\gamma^1_{\text{NZL frozen}}$							−0.012	0.002
							(0.012)	(0.002)
$\gamma^1_{\text{ROW frozen}}$								0.002
								(0.002)

Table S3. Estimated Coefficients from Demand System Regression for the Japanese Beef Import Demand Model (continued)

	USA Chilled Beef	AUS Chilled Beef	NZL Chilled Beef	ROW Chilled Beef	USA Frozen Beef	AUS Frozen Beef	NZL Frozen Beef	ROW Frozen Beef
$\gamma^2_{\text{USA chilled}}$	-0.586^{**}	0.538^{**}	0.035	0.013	-0.851^{***}	0.632^{***}	0.138^{*}	0.081
	(0.294)	(0.261)	(0.034)	(0.021)	(0.266)	(0.201)	(0.075)	(0.033)
$\gamma^2_{\text{AUS chilled}}$		-0.412^{*}	-0.124^{**}	-0.001	0.790^{***}	-0.549^{***}	-0.178^{**}	-0.064^{**}
		(0.222)	(0.052)	(0.020)	(0.245)	(0.175)	(0.083)	(0.030)
$\gamma^2_{\text{NZL chilled}}$			0.086^{***}	0.003	0.040^{**}	-0.046^{**}	0.017	-0.011^{***}
			(0.027)	0.005	(0.017)	(0.022)	(0.013)	(0.004)
$\gamma^2_{\text{ROW chilled}}$				-0.015	0.021	-0.038^{**}	0.022	-0.006
				(0.008)	(0.013)	(0.019)	(0.014)	(0.004)
$\gamma^2_{\text{USA frozen}}$					-0.529^{**}	0.392^{**}	0.129^{**}	0.008
					(0.212)	(0.166)	(0.055)	(0.020)
$\gamma^2_{\text{AUS frozen}}$						-0.07	-0.318^{***}	-0.005
						(0.117)	(0.095)	(0.020)
$\gamma^2_{\text{NZL frozen}}$							0.194^{***}	-0.006
							(0.060)	(0.010)
$\gamma^2_{\text{ROW frozen}}$								0.003
								(0.010)

Table S4. Estimated Coefficients from Demand System Regression for the Japanese Beef Import Demand Model (continued)

	USA Chilled Beef	AUS Chilled Beef	NZL Chilled Beef	ROW Chilled Beef	USA Frozen Beef	AUS Frozen Beef	NZL Frozen Beef	ROW Frozen Beef
$\gamma^3_{\text{USA chilled}}$	0.087	−0.085	−0.003	0.001	0.051	−0.061	0.01	0.000
	(0.056)	(0.056)	(0.006)	(0.006)	(0.041)	(0.038)	(0.019)	(0.008)
$\gamma^3_{\text{AUS chilled}}$		0.066	0.019**	−0.000	−0.054	0.042	0.014	−0.001
		(0.055)	(0.008)	(0.006)	(0.040)	(0.037)	(0.019)	(0.009)
$\gamma^3_{\text{NZL chilled}}$			−0.015	−0.001	0.012***	0.008*	−0.022***	0.001
			(0.010)	0.001	(0.004)	(0.005)	(0.003)	(0.000)
$\gamma^3_{\text{ROW chilled}}$				0.001	−0.009***	0.011	−0.002	−0.000
				(0.002)	(0.004)	(0.005)	(0.003)	(0.001)
$\gamma^3_{\text{USA frozen}}$					−0.022	−0.028	0.023*	0.027***
					(0.034)	(0.033)	(0.013)	(0.007)
$\gamma^3_{\text{AUS frozen}}$						0.058	−0.011	−0.018
						(0.038)	(0.018)	(0.007)
$\gamma^3_{\text{NZL frozen}}$							−0.004	−0.009**
							(0.014)	(0.003)
$\gamma^3_{\text{ROW frozen}}$								−0.000
								(0.003)

Notes: Estimates are derived from the three-regime smooth transition SDAIDS model with time-varying α,β and γ. The estimated transition function parameters are $\lambda_1 = \exp(\lambda_1^*) = \exp(-1.117)$, $c_1 = 0.077$, $\lambda_2 = \exp(\lambda_2^*) = \exp(1.131)$, $c_2 = 0.680$. Standard errors in parentheses. Single, double, and triple asterisks indicate significance at the $p < 0.10$, $p < 0.05$, $p < 0.01$ level, respectively.

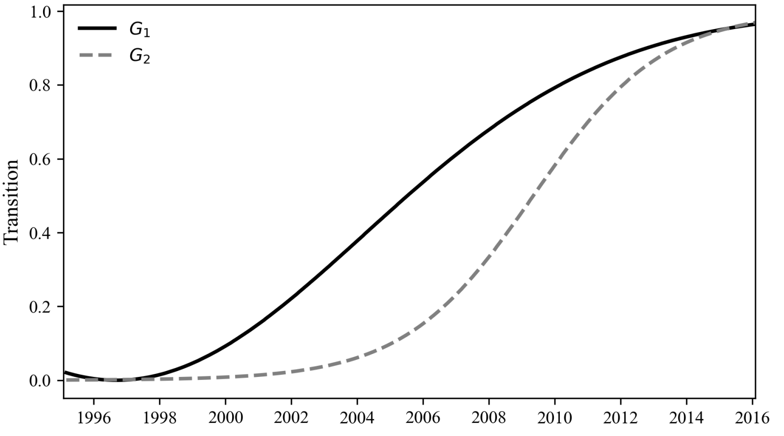

Figure S1. Estimated Transition Functions in the Japanese Beef Import Demand Model

Notes: Two smooth transition functions are used in the three-regime SDAIDS model estimation. The first is the exponential smooth transition function (labeled as G_1), with $\lambda_1 = \exp(-1.117) = 0.327$ and $c_1 = 0.077$. The second is the logistic smooth transition function (labeled as G_2), with $\lambda_2 = \exp(1.131) = 3.099$ and $c_2 = 0.68$. All four parameters are significant at the 1% level.

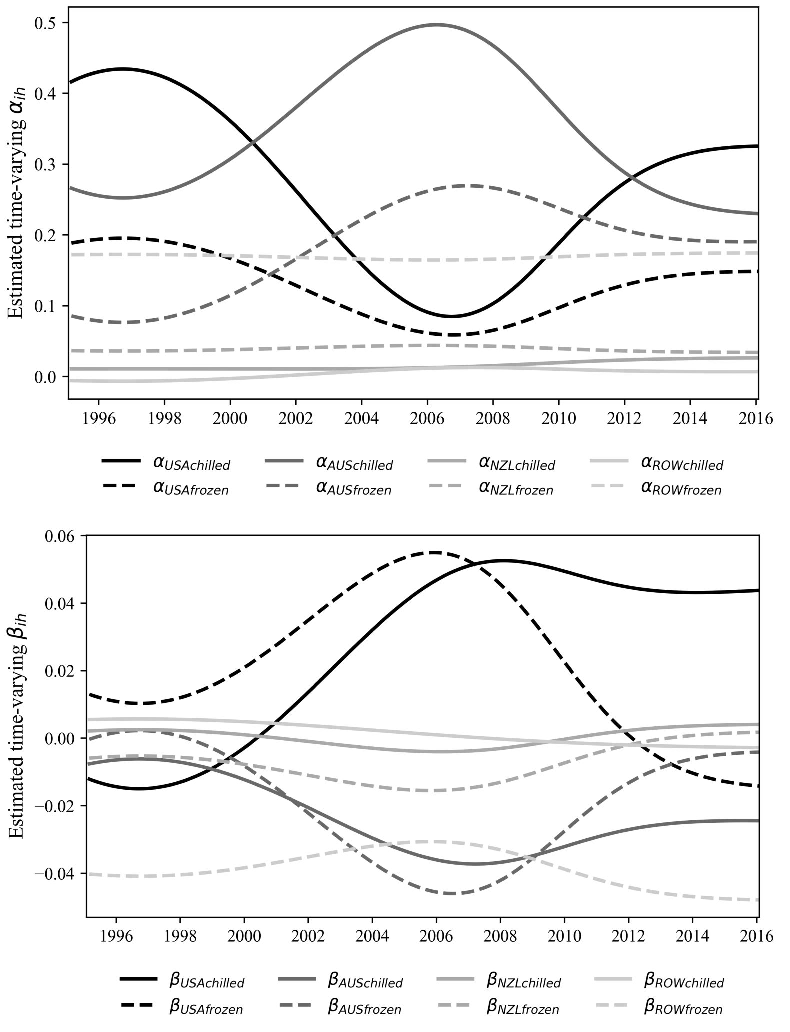

Figure S2. Estimated Time-Varying α_{ih} and β_{ih} in the Japanese Beef Import Demand Model

Notes: Time-series plots of α_{ih} and β_{ih}, $\forall i, h$ derived from the three-regime smooth transition SDAIDS model. For simplicity, time-series plots of γ_{ihjk} are estimated but not presented.