Impacts of Food Safety on U.S. Meat Demand

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and

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Abstract

We investigate the impacts of food safety on a weakly separable U.S. meat demand system (beef, pork, and poultry) using both the Generalized Almost Ideal Demand System (GAIDS) and the Rotterdam model. To measure food safety, indices are constructed based on the number of meat safety articles reported by the top 50 English language newspapers. The GAIDS permits estimation of food safety parameters in a theoretically consistent framework using the concept of demographic translation. The Rotterdam model offers a comparison of estimates to the GAIDS and a further test of the robustness of the food safety elasticities. We find that inferences with respect to food safety and autocorrelation are fragile to functional form choices. From the models investigated there is mixed evidence as to whether food safety concerns have impacted demand. Evidence from the GAI model indicates that food safety impacts could last for several quarters, whereas evidence from the Rotterdam model fails to reject the hypothesis that food safety variables are statistically different from zero over any period. There is also mixed evidence concerning autocorrelation. In the GAI model the problem of autocorrelation disappears by including food safety variables, which are found to be statistically significant and seemingly rectifying the misspecified model that omits food safety variables. This is not the case for the Rotterdam model where a correction for serial correlation is needed even in the presence of the food safety variables, which themselves are not statistically significant. The fragility of these inferences and estimated economic effects to specification choices, particularly to functional form and how the food safety variables enter the demand functions, make it difficult to draw many definitive conclusions about the magnitude or sign of food safety impacts on demand. Concerns of their statistical significance notwithstanding, the most definitive observation is that they are likely to be very small relative to price and expenditure effects and to other possible factors that may have impacted demand.

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Introduction

Food safety concerns in the United States (U.S.) have dramatically increased in the past decade with regard to incidences of contaminated meat products. Concerns have arisen because contaminated meat products can result in serious risk to the well being and health of consumers. Contaminated meat products come from a myriad of sources, including outbreaks of Listeria monocytogenes, Escherichia coli (E. coli), and Salmonella (Centers for Disease Control and Prevention). Food safety problems are not isolated to the U.S. Other unsafe contaminates in meats have emerged across the world, including highly publicized outbreaks of Bovine Spongiform Encephalopathy (BSE) in Europe. In this study, we attempt to investigate the effects of food safety on consumer's demand for meat (beef, pork, chicken and turkey) products in the U.S. over the past two decades.

The effects of non-price variables (i.e., demand shift or demographic variables) on aggregate meat demand in the U.S. have been studied extensively across competing consumer demand models.[1] McGuirk et al. (1995) augmented the intercept of the Linear Almost Ideal Demand System (LAIDS) with selected demographic variables and found that both health information and a changing labor force contributed to structural change in meat demand from 1960 to 1988. Brester and Schroeder (1995) and Kinnucan et al. (1998) examined the effects of advertising on U.S. meat demand with variations of the Rotterdam model. Both studies reported advertising impacts that were small, often not statistically significant, and unstable.[2] Piggott (1997) examined the demand response to generic advertising in the U.S. meat industry across a host of nested PIGLOG demand systems that incorporated demographic translation. The nested demand systems included, among others, the generalized translog and generalized AIDS, and found that advertising

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[1] The terms demographic variables, demand shifters, and shift variables are used interchangeably in this paper.

[2] See also Coulibaly and Brorsen (1999).

had indeed had a statistically significant effect on demand and that this finding was not sensitive to specific functional form used in estimation. LaFrance (1999) examined food consumption from 1918-1994 and incorporated age-distribution, ethnic background, and habit formation into his analysis, in addition to the traditional meat demand determinants. Using a quadratic model with demographic translation, LaFrance concluded ethnicity and age distribution both affected beef demand. Marsh, Schroeder, and Mintert (2000) examined the effects of health information, a changing labor force, and meat recalls on U.S. meat demand with the Rotterdam model. Health and labor force elasticities were similar, but not identical, to the magnitudes or signs reported by McGuirk et al. (1995). Meat recall elasticities were significant, but small. Evidence from the above studies illustrates that, while demand shift variables can be statistically significant, elasticity estimates are often small or not robust.

The fragile nature of elasticities of demand shift variables is neither unanticipated nor ignored in the literature.[3] For example, Deaton and Muellbauer have asserted that estimating price substitution elasticities alone is difficult in consumer demand models with time-series data. Wohlgenant found differences in demand elasticities for food across the Fourier flexible demand model, translog model, and generalized Leontief model. Piggott et al. (1996) examined the demand response to advertising in the Australian meat industry using the double log model, LAIDS, and the AIDS, correcting demand systems for autocorrelation. Elasticities estimates were found to be sensitive to autocorrelation and its specification. More recently, Alston, Chalfant, and Piggott have pointed out that demand shift variables must be incorporated in a manner that maintains the theoretical properties of the AIDS model, which ensures economic effects are invariant to the scaling of the data. They suggest including demand shift variables as modifications of pre-committed

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[3] Many meat demand studies have concluded that the impacts of competing meat prices on beef consumption are not stable. See, for example, Eales and Unnevehr, 1988; Moschini and Meilke, 1989; McGuirk et al., 1995. This finding suggests meat consumption patterns are determined by other factors in addition to relative prices and total meat expenditures.

quantities as in the generalized almost ideal system (GAIDS). Finally, Alston and Chalfant suggested that it is only prudent to investigate alternative specifications of consumer demand models when providing policy recommendations.

To date the impact of food safety on aggregate meat demand has received little attention in the agricultural economics literature. Several studies focusing on the impact of food safety on meat demand have targeted selected contaminants. For instance, Burton and Young, as well as Burton, Young, and Cromb, focused on the effects of food safety on meat demand in England using an index based on the number of newspaper articles generated about BSE. Meanwhile, Flake and Patterson focused on the effect of a single food safety index on meat demand in the U.S., which was constructed from the number of Associated Press articles on E. coli, salmonellosis, and BSE. Reported results suggest that effects of food safety on U.S. meat demand were modest and dominated by factors related to health information. These studies test the significance of food safety indices constructed from journal articles and the popular press to reflect consumer's information specific to the selected contaminant(s). As a result, inferences drawn from these results are specific to the events surrounding the selected contaminant(s).

Alternatively, Schroeder, Marsh, and Mintert used a linear AIDS model and Marsh, Schroeder, and Mintert used a Rotterdam model to quantify the impacts of beef, pork, and poultry product recall events on U.S. meat demand. Meat product recalls are relevant because they account for all listed contaminants reported by the USDA Food Safety Inspection Service. Consequently, inferences from models using meat product recalls may have more comprehensive implications to consumers, policy makers, and the meat industry. Both studies specified separate food recall indices for beef, pork, and poultry, estimated unconditional demand models, and reported similar but fragile estimates for meat recall parameters.

The primary objective of this paper is to empirically quantify the impacts of food safety on U.S. meat demand, incorporating food safety indices appropriately in a theoretically consistent consumer demand model. Meat types considered in this study are beef, pork, and poultry (chicken and turkey). Food safety indices are constructed separately for beef, pork, and poultry. The indices are based on the number of newspaper articles from the top 50 English language newspapers aggregated quarterly from 1980 to 1999. Consequently, food safety indices not only include information from meat recall events but also other issues such as BSE. Secondary objectives include examining the effect that alternative functional forms (e.g. Generalized Almost Ideal Demand System and Rotterdam demand models) that specify alternative methods to incorporate demand shift variables and autocorrelation have on food safety elasticity estimates.

Food Safety

Following earlier studies on meat safety (Burton and Young; Burton, Young, and Cromb; Flake and Patterson), food safety indices are constructed based on newspaper articles from the popular press. To enhance insight and analysis drawn from the information, food safety indices are constructed separately for beef, pork, and poultry.[4] Data for the series were obtained by searching the top fifty English language newspapers in circulation from 1980 to 1999, using the academic version of Lexis-Nexis search tool. The data were not weighted otherwise. Key words searched were *food safety* or *contamination* or *product recall* or *outbreak* or *salmonella* or *listeria* or *E. coli* or *trichinae* or *staphylococcus* or *foodborne*.[5] From this information base the search was narrowed to collect beef, pork, and poultry information separately by using additional terms a) beef and hamburger, b) pork and ham, and c) chicken, turkey, and poultry, respectively. The newspaper articles were then linearly aggregated quarterly to construct beef,

[4] Marsh, Schroeder, and Mintert demonstrated that a composite index made up of beef, pork, and poultry can confound food safety own- and cross-effects and unduly reduce the power of associated statistical tests.

[5] The list of contaminants were based on those listed in the USDA's Food Safety Inspection Services meat product recall data base. See http://www.fsis.usda.gov/OA/news/yrecalls.htm#RNR or Marsh, Schroeder, and Mintert.

pork, and poultry media indices.

Figure 1 shows the beef, pork, and poultry media numbers aggregated quarterly from 1980 to 1999. Beginning in 1980 the number of reported food safety articles for each series remained small, trending slowly upward until 1988. From 1988 and through 1999 the number of articles have moved upward sharply intertwined with dramatic peaks of information, dominated by the beef series. Over the study period the beef series exhibits the highest mean and most variation in the number of articles, with a mean of 153 articles and standard deviation of 223. Next is the poultry series that has a mean of 139 articles and standard deviation of 129. The pork series has a mean of 39 articles and standard deviation of 41. The maximum number of reported articles per quarter for beef is nearly 1200 in 1996, for poultry over 600 in 1999, and for pork nearly 200 in 1997. Not surprisingly, predominate peaks in the beef, poultry, and pork series relate to important events in the recent history of food safety.

Beef played the most prominent role in meat safety events as recorded by newspaper articles. In 1990 a Bovine Spongiform Encephalopathy (BSE) outbreak was reported in Europe that yielded an increase in food safety related articles in beef to 326 in the second quarter of the year. The 1993 peak of 432 articles in the first quarter of the year coincided with an isolated Escherichia coli (E. coli) outbreak in the state of Washington. In 1996, BSE news resurfaced after scientists in Europe linked BSE in beef to a new variant Creutzfeldt-Jakob disease (CJD) in humans. Nearly 1200 related articles were reported in the first quarter of 1996 alone. The 1997 peak in media reports was related to a massive recall of beef contaminated with *E. coli* that occurred in the midwest U.S. Other important events in the meat industry during the late 1990's included USDA's final rule on Pathogen Reduction/Hazard Analysis and Critical Control Point (PR/HACCP) systems. The PR/HACCP rule requires meat and poultry plants under Federal inspection to take responsibility for reducing the contamination of meat and poultry products with pathogenic bacteria.

Poultry has also played an important role in meat safety events reported by newspapers. Of the three series, poultry exhibited the first evident peak of media information in the fourth quarter of 1988. This was related to a salmonella outbreak in chickens and eggs resulting from providing chickens feed with animal remains. From 1980 through the third quarter of 1988 the average number of articles per quarter was 24. After the third quarter of 1988 and through the fourth quarter of 1999 the average number of articles per quarter sharply increased to 228. More recently a bird flu outbreak in poultry throughout Hong Kong and China lead to 571 newspaper articles in the last quarter of 1997.

Pork has contributed less to the number of media reports than either beef or poultry. The number of articles has steadily increased since 1980, but more slowly relative to the other series. From 1980 through the third quarter of 1988 the average number of articles per quarter was 9, while the average number of articles per quarter from the third quarter of 1988 to the fourth quarter of 1999 had increased to 62. The maximum number of articles for one quarter peaked at 241 in 1999, which followed a dioxin scare in pork in Europe. Nevertheless, meat safety issues in pork products remain important to consumers and industry as pork has been linked to outbreaks of listeria and other potentially dangerous contaminants.

An alternative way to interpret the media data is to consider the share of food safety articles reported about beef, pork, and poultry. Shares were created by dividing the beef, pork, and poultry indices by a composite index. The composite index was constructed by linearly aggregating the beef, pork, and poultry series into a single series. From 1980 to 1999 the average share value was 50%, 36%, and 14% for poultry, beef, and pork respectively. Up to the early 1990's poultry received on average the largest share of articles related to food safety with 55%, followed by 28% for beef and 17% for pork, respectively. Since the early 1990's beef's average share has increased to 44%, with the share of articles peaking at nearly 70% in 1997 and settling under 50% in 1999. Over the same period, poultry's average share has fallen to 44%, with the share of articles bottoming out at just over 20% in 1996 and settling under 40% in 1999. Relative to beef and poultry the share of pork has seen a smaller trend downward from 1980 to 1999, with the share of articles averaging 17% in the 1980's and 12% in the 1990's.

Demand Models

In this paper meat is treated as a weakly separable group in which consumption of an individual meat item depends only on the expenditure of the group, the prices of the goods within the group, and certain introduced demand shifters. This weakly separable group is comprised of three meats namely beef, pork, and poultry (chicken and turkey). Meat data used in the analysis are quarterly observations over the period 1982(1)-1999(3), providing a total of 71 observations (from various USDA databases). Food safety variables used in the analysis are quarterly data over the same period, consisting of the linearly aggregated beef, pork, and poultry indices discussed above. Finally, effects of time on meat demand are incorporated in the model through the use of quarterly demand shift (binary) variables for seasonality and a linear trend variable.

It is well known in empirical demand analysis that specification choices and functional form can influence inferences and estimates of economic effects. Since the auxiliary hypothesis of functional form is unavoidable one approach to take account of this influence is to estimate several functional forms and ascertain the robustness of inferences and estimated economic effects across alternative functional forms (Alston and Chalfant). The two most common approaches to estimating demand systems that incorporate demand shifters are the Almost Ideal Demand System (AIDS) (Deaton and Muellbauer) and Rotterdam model (Theil).[6] The AIDS model has been adopted extensively in the literature since is it is appropriate for

[6] A search of Econlit using key words "Almost Ideal Demand System" and "Rotterdam model" revealed than 156 papers and 43 papers, respectively.

aggregate and individual consumer analysis and allows restrictions from theory such as homogeneity, adding-up, and symmetry to be imposed. The Rotterdam model, which is derived from consumer demand theory, is a valid discrete approximation in variable space and is linear in parameters. Barnett and Mountain demonstrated that it is appropriate for aggregate and individual consumer analysis, respectively.

Often the applications of the AIDS model the sometimes difficult to estimate "true price index" is replaced with Stones price index. However, Moshinci (1995) demonstrated that Stone's Price index is not invariant to units of measurement rendering this approach problematic. More recently Alston, Chalfant and Piggott (2001) showed that the common approach of incorporating variables other than price and income into the AIDS model by augmenting the intercepts of the share equations, which was originally suggested by Deaton and Muellbauer (1980, p. 320), is also problematic. This is because estimates of economic effects (elasticities) are no longer invariant to units of measurement. Alston, Chalfant, and Piggott identified one practicable alternative, preserving other desirable features of the Almost Ideal model and also allowing demand shifters to be incorporated parsimoniously and flexibly, by adopting a generalized version of the Almost Ideal model, the Generalized Almost Ideal (GAI) model, first derived by Bollino. The expenditure function used to characterize the GAI model allows for a portion of total expenditures to be allocated to pre-committed quantities, which are unobservable and must be estimated along with the other parameters in the demand system. These pre-committed quantities can be modified to include demand shifters and simultaneously maintain the desirable theoretical properties of the model. The GAI model can, of course, be viewed as a generalization of the Linear Expenditure System (LES) in which the marginal budget shares are no longer constant but are instead of the Almost Ideal form.

For the purposes of this study we investigate the impact of the food safety variables on the demand for meat in the U.S. using the GAI model and the absolute price version of the Rotterdam model.

Estimating both functional forms is undertaken in an attempt to establish the robustness of inferences and estimated economic effects across models. The functional formulation for both models with definitions of parameters and variables can be written as:

Generalized Almost Ideal Demand Model:

$$w_i = \left(\frac{p_i c_i}{M}\right) + \left(\frac{M^*}{M}\right)\left(a_i + \sum_{j=1}^{N} g_{ij} \ln p_j + b_i \ln\left(\frac{M^*}{P}\right)\right) + \mathrm{v}_\mathrm{i}$$

where

$$M^* = M - \sum_{j=1}^{N} c_j p_j$$

$$\ln P = d + \sum_{\mathrm{j=1}}^{\mathrm{N}} a_k \ln p_j + \tfrac{1}{2} \sum_{\mathrm{j=1}}^{\mathrm{N}} \sum_{\mathrm{j=1}}^{\mathrm{N}} g_{kj} \ln p_k \ln p_j$$

Rotterdam Model (Absolute Price Version)

$$\overline{w}_i \Delta \ln x_i = \sum_{j=1}^{N} g_{ij} \Delta \ln p_j + b_i \Delta \ln \overline{q} + \mathrm{v}_\mathrm{i}$$

where

$$\overline{w}_{i,t} = \tfrac{1}{2}\left(w_{i,t} + w_{i,t-1}\right)$$

$$\ln \overline{q} = \sum_{k=1}^{N} \overline{w}_k \Delta \ln q_k$$

In these equation, q_i= per capita consumption of meat type i, p_i=per unit price of meat type i, M = total expenditures per capita on the different meat types ($M = \sum_{i=1}^{N} p_i q_i$), and w_i=expenditure share of meat type i ($w_i = \sum_{i=1}^{N} \frac{p_i q_i}{M}$). The coefficients a_i, g_{ij}, b_i, d, and c_j are parameters to be estimated and v_i is the random error term. General demand restrictions, which are derived from economic theory, can be imposed using parameter constraints with homogeneity being imposed by $\sum_{j=1}^{N} g_{ij} = 0$, adding up conditions by $\sum_{i=1}^{N} b_i = 0$ and $\sum_{i=1}^{N} a_i = 1$, and symmetry $g_{ij} = g_{ji} \quad \forall \ \mathrm{i} \neq \mathrm{j}$.

Incorporating Demand Shifters

To test the statistical significance of food safety proxied by media events we need to incorporate these variables in such a fashion so that each of the functional forms underlying consistency with theoretically is maintained. Furthermore, because the U.S. quarterly meat demand data seem to exhibit seasonal patterns and trends in consumption over time, it is also appropriate to include seasonal dummy and time trends variables into the above demand systems. Researchers need to proceed with caution in deciding how to incorporate demand shifters to avoid some not so obvious problems that can arise. For example, modifying the intercepts of the Almost Ideal demand system is a common approach but has the unfortunate implication that estimated economics effects (elasticities) are no longer invariant to units of measurement.

One way to avoid invariance problems in the AIDS is to adopt a generalized model that allows for pre-committed goods with demographic translation. The transformed expenditure function that includes pre-committed quantities yields demand functions that are made up of two components:

$$
\begin{aligned}
q_i &= c_i + q_i^* \left[\mathbf{p},\ M^*\right] \\
&= c_i + q_i^* \left[\mathbf{p}, \left(M - \Sigma\, p_i c_i\right)\right]
\end{aligned}
$$

where c_i is the pre-committed quantity and q_i^* is the supernumerary quantity of the i^{th} good. The distinction between the two types of consumption is important, since the pre-committed quantities are independent of prices and expenditure, whereas the supernumerary quantities are not. Given this distinction, it seems natural to augment the c_i's to be linear functions of demand shifters such as time and food safety variables. Note, however, that many researchers have ruled out this possibility by assuming that pre-committed consumption does not exist (for example, the usual AI make this assumption implicitly). It is also worthwhile to point out that augmenting the pre-committed quantities does not imply any restrictions on how any prospective demand shifters effects the demand for any particular good. The effect of a change from particular demand shifters can be positive or negative, depending on the relative magnitudes and signs of the direct and expenditure effects. The only required restriction is that the sum of changes in expenditures on pre-committed quantities must be equal and opposite to changes in supernumerary expenditures, leaving total expenditures unchanged. Thus, this translating approach is flexible in how the augmenting variables can affect the demand, and it is also parsimonious in terms of the additional parameters that must be estimated.

In the GAI model we modify the committed quantities, the 'c_i'*s*, to depend upon demand shifters in the following fashion:

$$\tilde{c}_j = c_{j0} + t_i t + \sum_{k=1}^{3} q_{ik} qd_k + \sum_{m=0}^{M} f_{i,m} bf_{t-m} + p_{i,m} pk_{t-m} + k_{i,m} py_{t-m}$$

where t is a linear time trend set equal to 1 in 1983:2 and qd_k (k=1, 2, and 3) are quarterly intercept dummies, bf_{t-m} is the beef food safety variable, pk_{t-m} is the pork food safety variable, and py_{t-m} is the poultry food safety variable all lagged m periods. In the Rotterdam model we include an intercept in the equation for each good to measure changes in consumption over time, seasonal dummy variables to capture seasonality, and food safety variables resulting in the following augmented model

$$\overline{w}_i \Delta \ln x_i = t_i + \sum_{k=1}^{3} q_{ik} qd_k + \sum_{m=0}^{M} f_{i,m} \Delta \ln bf_{t-m} + p_{i,m} \Delta \ln pk_{t-m} + k_{i,m} \Delta \ln py_{t-m} + \sum_{j=1}^{N} g_{ij} \Delta \ln p_j + b_i \Delta \ln \overline{q} + \mathrm{v}_i$$

In the above models the coefficients c_i, è$_i$, τ_i, ϕ_i, π_i, κ_i are additional parameters to be estimated.

An important point of contrast is the distinct differences of the functional relationships that specify how the food safety variables are enter into each model. In the GAI model the food safety variables enter in levels, while in the Rotterdam model the food safety variables enter as the logarithm of first differences.

Economic theory does not give us much guidance on how such demand shifters should enter except that it should be in a fashion that is flexible but preserves the theoretical properties of each of the models. Comparison of the resulting estimated economic effects across the GAI and Rotterdam models will serve to provide some evidence as to how appropriate and sensitive each alternative might be to modeling food safety demand shifters in the respective models.[7]

Autocorrelation

A final consideration is testing and correcting for autocorrelation in the meat demand models. Autocorrelation has been reported in aggregate U.S. meat data in recent studies using variations of both the AIDS and Rotterdam models (Marsh, Schroeder, and Mintert; Eales, Hyde, and Schrader). It has important implications with regard to specification choices, statistical inference, and estimated economic effects. Typically, if autocorrelation is found to be present in a demand system a correction is preformed that restricts the autocorrelation coefficient to be the same for every equation. We not only consider this case but also relax this constraint to allow coefficients to be different across equations and to allow the possibility of the cross-correlation from other equations.

More formally, to consider alternative forms of autocorrelation we follow Piggott et. al. (1996) and assume the vector of errors in the system of equations is determined by $\mathbf{e_t} = \mathbf{R}\mathbf{e_{t-1}} + \mathbf{v}_t$ for t = 2, …., T, where $\mathbf{v}_t$'s are independent $N(0, \Sigma)$ random vectors, and $\mathbf{R}$ is an $n \times n$ matrix of unknown parameters. Berndt and Savin showed that when $\mathbf{e_{t-1}}$ and $\mathbf{v_t}$ are statistically independent, the adding up property of the shares ($i'\mathbf{w}_t = 1$ where $\mathbf{i}$ is a $n \times 1$ vector of ones and $\mathbf{w_t}$ is the vector shares) implies a restriction $i'\mathbf{R} = k$, where k is an unknown constant. This restriction $i'\mathbf{R} = k$, Berndt and Savin showed can be

7 One empirical concern is the extreme changes in magnitude of the food safety variables from quarter-to-quarter, especially beef. In the GAI model, by incorporating level variables as demand shifters, the impact is direct and not

transformed into the more tractable form of $i'\overline{\mathbf{R}} = 0$, where $\overline{\mathbf{R}}$ is an $n \times n-1$ matrix with elements $\overline{R}_{ij} = R_{ij} - R_{in}$ for i=1, …, n and j= 1, …, n-1. In practice to implement this general form of autocorrelation correction we estimate $\overline{\mathbf{R}}^*$ which is the $n-1 \times n-1$ matrix formed by the first $n-1$ rows of the $\overline{\mathbf{R}}$. That is, it is the elements of $\overline{\mathbf{R}}^*$ that are estimated not $\overline{\mathbf{R}}$ or **R**. However, the estimates of $\overline{\mathbf{R}}^*$ combined with above mentioned restrictions imposed on it's elements and any zero restrictions can be used to recover the estimates of **R** if needed. Generally, actually solving for the individual R_{ij}'s is not as important as simply knowing whether they are jointly statistically significant which can be established from testing the statistically significance of $\overline{\mathbf{R}}^*$.

Model Results

In the empirical application three alternative autocorrelation corrections are considered: (a) a null **R** matrix (N-R^{matrix}) with all elements restricted to zero, specifying no autocorrelation; (b) a diagonal **R** matrix (D-R^{matrix}) wherein all diagonal elements are restricted to be identical and all off-diagonal elements are restricted to zero; and (c) a full **R** matrix (F-R^{matrix}) where all elements of **R** matrix are non-zero. Columns 1, 2, and 3 in Table 2 provide estimates of coefficients, standard errors, and summary statistics from the GAI model with alternative autocorrelation corrections without food safety variables. Columns 1, 2, and 3 in Table 3 report similar estimates for the Rotterdam model with alternative autocorrelation corrections without food safety variables. The estimates in Table 2 reveal positive and statistically significant pre-committed quantities of beef (c_{10}'s). In contrast, pre-committed quantities for pork and poultry (negative) are not statistically significantly different from zero. In addition, the estimated coefficients indicate statistically significant effects of seasonality and time trends in the pre-committed quantities. Findings of statistically significant seasonality and time trends are also supported by estimates for the Rotterdam model

dampened. In the Rotterdam model the natural log transformation dampens this impact and provides diminishing marginal in Table 3. Thus the variables representing seasonality and time trends are retained as a maintained hypothesis.

Table 2 (columns 4, 5, and 6) shows coefficient estimates from the GAI model combining alternative autocorrelation corrections with food safety variables. Similarly, Table 3 (columns 4, 5, and 6) reports estimates using the Rotterdam model coupling autocorrelation corrections with food safety variables. The estimates in Table 2 reveal positive and statistically significant pre-committed quantities of beef (c_{10}'s), pork (c_{20}'s) and poultry (c_{30}'s) after the food safety variables have been incorporated into the models. The individual coefficient estimates for the seasonal (θ_k's) and trend variables (τ_i's) remain statistically significantly different from zero in both the GAI and Rotterdam models across the alternative autocorrelation corrections, further supporting the decision to include these demand shifters into the models as a maintained hypothesis.

Turning to the impact of food safety on meat consumption, both contemporaneous and lagged food safety variables are included in the model specification to allow for possible dynamic effects. Previous studies indicate significant carryover effects of information between contemporaneous and lagged food safety variables (Marsh, Schroeder, and Mintert). That is, the effects of a given "media event" as capture by the food safety indexes may be spread over time. A priori there is no way of knowing how long effects on consumption of a given dose of a "media event" will last (i.e., the length of the lag, m)—this is an empirical question. In the above specifications, the total food safety effect in each equation is determined by a moving average of current and lagged food safety variables with the weights estimated econometrically. This total effect in each equation for each type of food safety information will be referred to as the "stock of food safety."

To investigate the carryover effect of food safety information both the GAI and Rotterdam models were estimated with the respective stocks of food safety variables in each equation having lag lengths varying from zero (m=0, the current level of food safety) to four quarters (m=4, the current value plus four quarters of lagged values). No further lags were considered, since each additional lag requires an additional nine parameters (three goods and three separate food safety variable) to be estimated, and longer lags

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effects to an additional newspaper article.

created convergence problems in the GAI model. The Rotterdam model is more parsimonious in the number of additional parameters that are required for additional lags only requiring an additional six parameters.[8] Table 4a shows results of likelihood ratio tests for the GAI model, testing the joint statistical significance of incorporating an additional lag for the food safety variables.[9] The results support the inclusion of the current food safety variables (m=0) as well as the first period lagged (m=1) and third period lagged (m=3). Neither the addition of m=2 or m=4 were jointly statistically significant, which was robust across the alternative choices for autocorrelation corrections. Based on these findings we conclude the appropriate lag choice for the GAI model was m=3. Hence, it is the estimated coefficients from the GAI model (with m=3) and the alternative autocorrelation corrections that are reported in Table 2.

The above results are not robust across the Rotterdam and GAI models. For the Rotterdam model, likelihood ratio tests are also used to test the statistical significance of addition lags of food safety variables up to four periods lagged (m=4). Table 4c reveals that none of the food safety variables m=0, 1, 2, 3, or 4 are joint statistically significantly different from zero. Despite the lack of statistical significance of the additional lags of food safety we chose to report m=3 for the Rotterdam model in Table 3 to compare results across different functional forms. A variety of factors are likely contributing to the differences in likelihood ratio tests across models. One is that the GAI model accounts for pre-committed goods while the Rotterdam model does not. If this specification is in fact correct, which is consistent with economic theory and is supported by the statistical results, then the GAI model should better delineate the effects of food safety variables and the Rotterdam model may be misspecified and not expected to provide correct inferences.

Hypothesis Tests of Model Specification

Assuming the appropriate lag structure of food safety variables is m=3, we test the joint significance of the food safety variables across the alternative autocorrelation specifications discussed above. Results are

[8] The weights on the lagged values of advertising were estimated as unconstrained parameters. Placing restrictions on the lag weights is a way of reducing the number of parameters to be estimated. For example, Piggott et al. (1996) required lag weights be the same across equations for a given type of advertising using Australian data. Alternatively, Brester and Schroeder imposed a geometric lag structure on advertising using US data.

reported for the GAI and Rotterdam models in Tables 4a and 4c, respectively. Looking across alternative autocorrelation specifications, results reconfirm for the GAI model the finding that food safety variables are jointly statistically significantly different from zero. Similarly, the above results for the addition of individual lags are echoed in the Rotterdam model with Table 4c, confirming that the food safety variables are not jointly statistically significant from zero across alternative autocorrelation corrections. The Rotterdam results obviously conflict with the findings from the GAI model and suggests that inferences with respect to whether food safety variables maybe fragile.

Next we consider the issue of alternative autocorrelation corrections. Table 4b tabulates results that test the three forms of autocorrelation corrections that are estimated for all of the alternative lags of food safety variables. The results reveal that when no food safety variables are included, which have been determined to be statistically significantly different from zero, autocorrelation cannot be rejected. The null of no autocorrelation (N-R^{matrix}) is rejected against the both the diagonal (D-R^{matrix}) and general autocorrelation (F-R^{matrix}) specifications. Hence, this version of the GAI model appears to be misspecified. Table 4b also reveals that upon the inclusion of at least the current level of food safety variables, and further additional lagged food safety effects, no autocorrelation correction is required. It appears that the omitted food safety variables may mask, or be masked by, the autocorrelation problem. The lack of significant autocorrelation provides some confidence for the inclusion of the models that include the food safety variables. Finally, we never reject the diagonal autocorrelation correction (D-R^{matrix}) against the alternative general autocorrelation correction (F-R^{matrix}).

Table 4d tabulates the hypothesis tests for the various autocorrelation corrections estimated over all of the alternative lags of food safety variables considered for the Rotterdam model. The null hypothesis of

9 The likelihood ratio statistic reported in Table 4 is an adjusted likelihood ratio appropriate for systems of equations (see Bohm, Rieder, and Tinter; Bewley). Although not reported here, the adjusted likelihood ratio test suggested by Moschini, Moro, and Green yielded results that were consistent with inference drawn here.

no autocorrealtion (N-R^{matrix}) is rejected against the alternative of a diagonal R matrix (D-R^{matrix}) for all lag lengths of food safety considered. It appears for the Rotterdam model that inclusion of the food safety variables does not curtail the presence of serial correlation in the error terms and that at least a single coefficient for each equation (D-R^{matrix}) is required. Similar to the GAI results, we never rule in favor of the general autocorrelation correction (F-R^{matrix}) over the null of a diagonal autocorrelation correction (D-R^{matrix}).

To summarize the results of the above inferences the results appear fragile and sensitive to the maintained hypothesis of functional form and to the specification choices of how food safety enters the demand system. However, given either the GAI or Rotterdam model, the results are quite robust to choices of autocorrelation corrections. Under the maintained hypothesis that the GAI model is the correct functional form, we find that the food safety variables that augment the committed quantities are statistically significantly different from zero and that there are important dynamics from these effects that last three-quarters. This finding is robust across alternative autocorrelation corrections, which themselves are insignificant once the food safety variables are incorporated into the model. Alternatively, under the maintained hypothesis that the Rotterdam model is the correct functional form, we find that the food safety variables are not statistically significantly different from zero and as a result that there are no important dynamic carryover effects. This finding is robust across alternative autocorrelation corrections with a single coefficient autocorrelation correction (D-R^{matrix}) seeming to be sufficient for the Rotterdam model.

Estimated Economics Effects

Table 5 provides estimates of means of the Marshallian price elasticities, expenditure elasticities, and food safety elasticities that are calculated at every data point. Elasticities reported are for both models shown in Tables 2 and 3 with food safety variables included for alternative specifications of autocorrelation corrections. For the food safety elasticities both immediate-effect and total-effect responses are provided. The immediate effects are the short-run elasticities of demand response to food safety measuring the percentage change in consumption of the ith good in response to a one-percent increase in the kth type of food safety variable $z_{k,t}$ (i.e., $w_{ik} = \partial \ln q_{i,t} / \partial \ln z_{k,t}$). The total effects are the demand response to food safety that measure the percentage change in consumption of the ith good in response to a one-percent permanent increase in the kth type of food safety variable $z_{k,t}$ (i.e., $m_{ik} = \sum_{m=0}^{3} \partial \ln q_{i,t} / \partial \ln z_{k,t-m}$).

A striking feature of the price and expenditure elasticities is that they are remarkable robust across the alternative autocorrelation specifications for each of the functional forms. That is, the differences in the estimates of the price and expenditure elasticities appear to be somewhat consistent across functional forms. For example, the own-price elasticities of demand are consistently estimated to be more elastic (only modestly) using the GAI model than the Rotterdam model. The Marshallian cross-price elasticities of demand are negative across both models indicating that these meats are gross complements within this weakly separable group with one exception in the GAI model—the cross-price elasticity of demand for poultry with respect to beef price. The cross-price elasticity's of demand for pork and poultry are predominately larger in magnitude for the GAI model when compared to their Rotterdam counterparts, signifying a larger complementary effect. This is not true for beef with the cross-price elasticities smaller using the GAI model.

In relation to consistency with economic theory, both models perform favorably. The Rotterdam model slightly outperforms the GAI model with the curvature requirements of negative semi-definiteness being satisfied globally within the sample across the variations of autocorrelation corrections. The GAI satisfied the requirements with slightly more than 90 percent of the sample. The difference in consistency with theory might come about from the parsimony in terms of number of parameters that are required to be estimated in the Rotterdam with demand shifters entering as modifications of the intercepts compared to modification of committed quantities in the case of the GAI model.

Turning to the focus of the paper, which is the estimated economic effects of the food safety variables, we see much less similarity in the results across functional forms. But once again the results seem quite robust across choices of autocorrelation corrections, at least for the GAI model. There are more differences in the estimated food safety elasticities across choices of autocorrelation for the Rotterdam model when we compare the estimates from N-R^{matrix} with the other specifications. Recall that the N-R^{matrix} was consistently rejected against the alternative hypothesis of D-R^{matrix} implying that properly accounting for autocorrelation in this model appears as though it can have important impacts on the magnitude of estimated effects (their lack of statistical significance notwithstanding). For example, the total effect on the own-beef food safety variables using the N-R^{matrix} was –0.00293 compared to estimates of -0.00491 and –0.00499 using the N-D^{matrix} and F-R^{matrix}, respectively, a difference in magnitude of more than 150 percent. A similar comparison can be illustrated for the estimates for pork. With these autocorrelation differences noted the remainder of discussion will then focus on the differences in estimates across functional form choices.

A priori, we expect the own-food safety variables be negative (i.e., $w_{ii} < 0$ and $m_{\text{ii}} < 0$). That is, we expect that increases in the media index reflecting increases in incidents of food safety events or concerns would adversely affect the demand for the particular meat. The most striking feature of the estimated food safety elasticities is that the magnitudes involved are very small compared to those of prices and expenditure, irrespective of functional form. Inspecting the immediate short-run own-food safety elasticities based on the GAI model, it's apparent they do not conform well to prior beliefs. For example, the own-food safety elasticity for beef (w_{11}) and pork are almost always positive across the entire sample. Encouragingly, things improve for the estimate of the own-food safety elasticity for poultry with this estimate being negative for more than 90% of the sample and managing to generate a mean that indeed is of the expected sign and around –0.014. The total-effect own-food safety elasticities generated a larger proportion of the sample with estimates for beef (m_{11}) and pork (m_{22}) that confirm with priors, around 34 percent for beef and 18 percent for pork. For poultry (m_{33}) less of the sample conforms, reduced to around 80 percent compared to more than 90 percent for the short-run. Thus, despite finding statistically significant impacts from food safety variables in the GAI model and a lack of serial correlation, this model generates food safety elasticities that are difficult to interpret. However, it does appear that attempting to capture possible dynamic effects from food safety results in a larger proportion of the estimated economic effects becoming consistent with priors in relation to sign.

Similarly, for the Rotterdam model, we find the immediate short run own-food safety elasticities to be globally inconsistent with prior beliefs. The exception being the own-pork food safety variable in the model with no autocorrelation correction (N-R^{matrix}), but a model that we also reject. However, a more encouraging result occurs for the total effect own-food safety elasticities for beef (m_{11}) and pork (m_{22}), which are globally consistent with priors being negative across all three variations of autocorrelation corrections. The means of these elasticities from the preferred D-R^{matrix} model are –0.00491 and –0.00650, respectively. The total-effect estimate for poultry is estimated to be positive globally across the alternative autocorrelation corrections for poultry (m_{33}).

The immediate-effect and total-effect cross-elasticities from the food safety variables mostly are not consistent in sign across the GAI and Rotterdam models reported in Table 5. Exceptions where signs of cross-elasticities have the same sign across models include w_{21} and w_{13} (negative) and w_{23} (positive) for the short-run and m_{31} (positive) for the total-effects. There is much less variation in sign and magnitudes of the cross-estimates for the food safety variables within functional forms and across alternative autocorrelation corrections. A priori we would expect that these cross-elasticities might be positive anticipating that an increase in food safety concerns about a particular meat might induce an increase in demand for another meat. With only one of the cross-elasticities estimated to be positive across both models however, the empirical evidence does not provide much support for this prior belief. Interestingly, the GAI model estimates of the cross-elasticites tend to be positive more often than the Rotterdam model for both the immediate and long-run effects. In all, the results are mixed and inconclusive as to whether a specific cross-effect has a positive or negative impact on beef, pork, or poultry demand, hinging once again on specification choices of functional form and how food safety variables enter. These, inconclusive inferences drawn from the estimated cross-elasticities are not so surprising given a similar finding for the own-elasticities for the food safety variables. The lack of agreement across functional forms once again highlights the fragility of estimates to functional form choices

Overall, these very different results across the two functional forms highlight the apparent fragility of not only the inferences concerning whether the food safety variables are statistically significant or not —the GAI model inferences reject these variables are zero while the Rotterdam model fails to reject the null hypothesis that they are zero—but also the sign of the estimated effects of the food safety variables with respect to the choice of specification of functional form and how the food safety variables are model as affecting demand. Neither the inferences concerning statistical significance nor the estimated economic effects seem as sensitive to specification choices concerning autocorrelation.

Further Discussion

To better illustrate the underlying differences in the elasticity estimates from the GAI and Rotterdam models, plots of each of the total-effect own-food safety elasticities for the GAI and Rotterdam model are provided in Figures 2, 3, and 4. These are helpful in highlighting the fragility of estimates across functional form and the apparent differences of incorporating the food safety variables as modifications of the committed quantities compared to demand shifters in the Rotterdam models. In each of these figures the

GAI model generates elasticities that vary quite significantly across the sample and appear to be much more volatile relative to the Rotterdam elasticities. In particular, the plots reveal the fragility of estimates to specification choices. For instance, when the estimates for m_{11} and m_{22} conform to priors using the Rotterdam model, the GAI model generates estimates that are positive for the majority of the sample. Moreover, when the GAI model finds most of the estimates for poultry (m_{33}) to be negative the Rotterdam model estimates this elasticity to be positive globally. Finally, the increased volatility of the GAI estimates coincides with the increase in food safety articles (and associated volatility) in the late 1990's. In contrast the Rotterdam elasticities appear relatively stable over the entire sample.

This scenario makes it difficult to draw definitive conclusions about the impact on food safety on U.S. meat demand except to make the observations that the estimated economic effects are likely to be small in comparison to price and expenditure effects, sensitive to specification choices (particularly that of functional form and how the food safety variables are incorporated in the demand model), and inconclusive to whether or not the effects of a given dose of food safety information has significant carryover effects lasting several quarters. This perceived lack of consistency of elasticities across demand models and/or the inconsistency with a priori expectations may have arisen for several reasons. First, the food safety indices are specified only as linear aggregations of the number of articles per quarter from the top 50 English language newspapers. As noted the articles were not otherwise weighted. One possible alternative is to weight the articles by circulation of each newspaper. Other alternatives include using other media information or classifying articles as either positive or negative to construct a net index of food safety. Second, we have made no attempt to delineate the impacts of non-domestic information, say related to BSE in Europe, from information of contaminants originating within the U.S. Given the data can be partitioned, this result could be tested. Third, we imposed no parameter or functional restrictions on the lagged food safety variables. Freely estimating the lagged parameters in this manner may result in an over parametization of the underlying impact making it difficult to get precise estimates of the true impact of these food safety effects. A more restricted and parsimonious specification of how food safety variables may enter the demand system might be more appropriate and remains as a topic for further investigation.

Conclusion

Food safety concerns by consumers in the U.S. have dramatically increased in the past decade with regard to incidences of contaminated meat products. To date the impact of food safety on aggregate meat demand has received little attention in the agricultural economics literature. Hence, the primary objective of this paper is to empirically quantify the impacts of food safety on U.S. meat demand, incorporating food safety indices appropriately in a theoretically consistent consumer demand model. Meat types considered were beef, pork, and poultry (chicken and turkey). Food safety indices were constructed separately for beef, pork, and poultry. The indices are based on the number of newspaper articles from the top 50 English language newspapers aggregated quarterly from 1980 to 1999. Consequently, food safety indices not only include information from meat recall events but also other issues such as BSE.

Applying both the Generalized Almost Ideal Demand System (GAIDS) and the Rotterdam model, we tested the statistical significance of the food safety indices on demand for meat in the U.S. Using the GAI model we found statistically significant effects from food safety variables that last three quarters after the initial quarter. This finding was robust across alternative autocorrelation corrections, which were apparently not necessary as the GAI models with the food safety variables did not suffer from serial correlation. One of the most striking features of the estimated food safety elasticities is their small magnitude relative to price and expenditure elasticities. In all, the estimated economic effects of the food safety variables from this model were difficult to rationalize with a large proportion of the own-food safety elasticities for beef and pork being of the unexpected sign (positive) across the sample. This lack of consistency with priors across the sample may be a due to some misspecification on how food safety variables were specified or to some omitted variables that may have impacted demand such as advertising, habits, and other health effects. If the true underlying effects are indeed very small, then in practice it may be very difficult to disentangle these effects from the many other factors that may have impacted demand consistently over the entire sample.

To check the robustness of these findings with the GAI model we also estimated the Rotterdam model and incorporated the food safety indices into the model as demand shifters. The comparison served to highlight the fragility of inferences and estimated economic effects to specification choices. Using the Rotterdam model the food safety indices were consistently found to be insignificant for any specified lag length. Furthermore, the residuals for the Rotterdam model suffered from serial corrrelation even in the presence of the food safety variables (when the GAI specification did not). Finally, for the one total-effect own-food safety elasticity that the GAI model generated that conformed with the priors, namely poultry (m_{33}), the Rotterdam model estimate did not conform globally. Moreover, for the total-effect own-food safety elasticities estimated with the GAI model that did not conform to the priors namely beef (m_{11}) and pork (m_{22}), the Rotterdam estimates conformed globally.

Overall from the mixed evidence found in these alternative specifications, the only definitive conclusion that we can draw is that the impacts that food safety if they have indeed impacted demand at all, then are likely to be small. Any other conclusions about magnitudes of impacts and even sign at least from the alternative models investigated here hinge on specification choices. In reality because these food safety effects appear to be very small they are likely going to be difficult to disentangle from all of the other factors that may have more significantly impacted demand over the last several decades.

Table 1. Summary Statistics of Quarterly Data, 1982(1)-1999(3).

	Average	Std. Dev.	Minimum	Maximum
Beef Consumption (lbs./capita)	17.799	1.353	15.892	20.818
Pork Consumption (lbs./capita)	12.789	0.685	11.562	14.492
Poultry Consumption (lbs./capita)	19.607	3.040	13.674	24.767
Retail Beef Price (cents/lb.)	263.785	23.975	222.733	300.400
Retail Pork Price (cents/lb.)	206.676	24.138	167.800	248.100
Retail Poultry Price (cents/lb.)	90.068	8.636	72.103	105.121
Meat Expenditure ($/capita)	90.951	8.316	75.660	108.436
Beef Expenditure Share	0.516	0.038	0.435	0.586
Pork Expenditure Share	0.290	0.014	0.265	0.323
Poultry Expenditure Share	0.194	0.030	0.133	0.243
Beef Food Safety	162.817	223.358	2.000	1158.000
Pork Food Safety	41.887	40.925	0.000	241.000
Poultry Food Safety	151.296	126.822	6.000	571.000

Table 2: Estimated Coefficients for the Generalized Almost Ideal Model With and Without Food Safety Variables

	No Food Safety			With Food Safety		
	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}
LL	*548.790*	*558.558*	*561.140*	*626.008*	*626.370*	*629.457*
d	8.093*	6.557*	7.146*	22.657	23.425	19.191
	(2.031)	(1.864)	(2.115)	(15.614)	(15.742)	(13.718)
a_1	3.002*	1.534	2.112	12.444	12.982	10.917
	(1.286)	(1.048)	(1.276)	(7.893)	(8.016)	(7.618)
a_2	-0.975	-0.227	-0.456	-6.160	-6.351	-5.290
	(0.625)	(0.370)	(0.488)	(4.290)	(4.303)	(4.076)
g_{11}	3.809*	1.410	2.369*	12.719*	13.556*	12.956*
	(1.128)	(0.796)	(1.053)	(4.084)	(4.194)	(4.363)
g_{12}	-1.622*	-0.369	-0.681	-6.676*	-6.997*	-6.697*
	(0.672)	(0.351)	(0.529)	(2.364)	(2.398)	(2.478)
g_{22}	0.864*	0.249	0.328	3.584*	3.687*	3.553*
	(0.372)	(0.139)	(0.236)	(1.355)	(1.360)	(1.383)
b_1	0.757*	0.622*	0.686*	0.675*	0.681*	0.738*
	(0.112)	(0.132)	(0.135)	(0.145)	(0.142)	(0.167)
b_2	-0.345*	-0.211*	-0.238*	-0.363*	-0.361*	-0.391*
	(0.065)	(0.065)	(0.078)	(0.081)	(0.077)	(0.089)
c_{10}	17.229*	13.862*	15.559*	15.952*	16.100*	16.498*
	(1.720)	(3.068)	(2.359)	(1.228)	(1.158)	(1.391)
c_{20}	-3.094	-5.285	-0.687	3.956*	4.346*	4.247*
	(3.734)	(5.820)	(4.247)	(1.839)	(1.666)	(1.957)
c_{30}	-6.714	-19.995	-12.703	7.985*	8.192*	7.324*
	(7.309)	(16.080)	(11.404)	(2.832)	(2.670)	(3.199)
q_1	0.104	0.033	0.039	0.114	0.110	0.135
	(0.141)	(0.151)	(0.145)	(0.072)	(0.071)	(0.069)
q_2	0.818*	0.752*	0.767*	1.084*	1.073*	1.096*
	(0.138)	(0.166)	(0.152)	(0.089)	(0.085)	(0.083)
q_3	0.966*	0.947*	0.993*	1.290*	1.267*	1.258*
	(0.139)	(0.145)	(0.142)	(0.082)	(0.081)	(0.083)
t_1	0.056*	0.095*	0.067*	0.039*	0.036*	0.034*
	(0.014)	(0.031)	(0.021)	(0.008)	(0.007)	(0.007)
q_4	-1.067*	-1.106*	-1.106*	-1.213*	-1.213*	-1.208*
	(0.148)	(0.129)	(0.115)	(0.080)	(0.081)	(0.075)
q_5	-1.440*	-1.478*	-1.470*	-1.427*	-1.429*	-1.424*
	(0.150)	(0.150)	(0.128)	(0.095)	(0.092)	(0.087)
q_6	-1.109*	-1.115*	-1.077*	-0.913*	-0.931*	-0.930*
	(0.141)	(0.120)	(0.110)	(0.081)	(0.082)	(0.078)
t_2	0.099*	0.119*	0.090*	0.082*	0.079*	0.074*
	(0.020)	(0.031)	(0.022)	(0.011)	(0.011)	(0.011)
q_7	-2.418*	-2.416*	-2.444*	-2.566*	-2.579*	-2.568*
	(0.142)	(0.141)	(0.136)	(0.085)	(0.088)	(0.081)
q_8	-1.668*	-1.654*	-1.686*	-1.945*	-1.956*	-1.933*
	(0.145)	(0.168)	(0.153)	(0.095)	(0.095)	(0.095)
q_9	-1.252*	-1.265*	-1.223*	-1.342*	-1.342*	-1.335*
	(0.137)	(0.130)	(0.131)	(0.091)	(0.094)	(0.087)
t_3	0.185*	0.234*	0.209*	0.172*	0.171*	0.171*
	(0.022)	(0.042)	(0.030)	(0.010)	(0.009)	(0.010)

Continued ……

Table 2: Estimated Coefficients for the Generalized Almost Ideal Model With and Without Food Safety Variables (Continued ……)

	No Food Safety			With Food Safety		
	N–R matrix	D–R matrix	F–R matrix	N–R matrix	D–R matrix	F–R matrix
Beef Food Safety Variables						
f_{10}				2.29E-04	2.57E-04	1.67E-04
				(4.28E-04)	(4.04E-04)	(3.83E-04)
f_{20}				-2.18E-03*	-2.18E-03*	-2.21E-03*
				(4.17E-04)	(4.01E-04)	(3.72E-04)
f_{30}				-8.00E-04*	-7.70E-04*	-9.90E-04*
				(3.09E-04)	(3.06E-04)	(3.10E-04)
f_{11}				-3.13E-03*	-3.08E-03*	-3.12E-03*
				(4.64E-04)	(4.59E-04)	(4.65E-04)
f_{21}				-8.20E-04	-7.90E-04	-6.80E-04
				(4.52E-04)	(4.55E-04)	(4.25E-04)
f_{31}				1.35E-03*	1.36E-03*	1.50E-03*
				(4.47E-04)	(4.56E-04)	(4.38E-04)
f_{12}				1.19E-03*	1.22E-03*	1.30E-03*
				(5.59E-04)	(5.31E-04)	(5.27E-04)
f_{22}				-8.20E-04	-7.60E-04	-8.70E-04
				(7.26E-04)	(7.08E-04)	(6.44E-04)
f_{32}				-1.22E-03*	-1.25E-03*	-1.29E-03*
				(5.19E-04)	(5.30E-04)	(4.96E-04)
f_{13}				-1.23E-03*	-1.23E-03*	-1.33E-03*
				(3.72E-04)	(3.56E-04)	(3.33E-04)
f_{23}				-1.57E-03*	-1.62E-03*	-1.40E-03*
				(3.04E-04)	(2.95E-04)	(2.82E-04)
f_{33}				-1.29E-03*	-1.24E-03*	-1.31E-03*
				(3.11E-04)	(3.13E-04)	(3.17E-04)
Pork Food Safety Variables						
p_{10}				-4.63E-03	-4.24E-03	-5.45E-03*
				(2.81E-03)	(2.69E-03)	(2.60E-03)
p_{20}				-6.30E-04	-3.80E-04	-6.20E-04
				(2.47E-03)	(2.42E-03)	(2.32E-03)
p_{30}				5.61E-03*	5.94E-03*	5.75E-03*
				(1.99E-03)	(1.99E-03)	(1.98E-03)
p_{11}				-1.45E-02*	-1.40E-02*	-1.42E-02*
				(4.55E-03)	(4.39E-03)	(4.17E-03)
p_{21}				-2.42E-02*	-2.37E-02*	-2.36E-02*
				(3.56E-03)	(3.38E-03)	(3.19E-03)
p_{31}				-1.64E-02*	-1.66E-02*	-1.78E-02*
				(3.80E-03)	(3.79E-03)	(3.84E-03)
p_{12}				2.96E-03	2.88E-03	3.06E-03
				(3.92E-03)	(3.75E-03)	(3.57E-03)
p_{22}				3.32E-03	3.37E-03	3.64E-03
				(3.84E-03)	(3.75E-03)	(3.50E-03)
p_{32}				6.86E-04	1.01E-03	1.29E-03
				(3.05E-03)	(3.13E-03)	(3.05E-03)
p_{13}				-4.61E-03	-5.11E-03	-4.56E-03
				(3.45E-03)	(3.31E-03)	(3.27E-03)
p_{23}				1.87E-02*	1.86E-02*	1.75E-02*
				(4.11E-03)	(4.08E-03)	(3.95E-03)
p_{33}				1.16E-02*	1.21E-02*	1.24E-02*
				(3.97E-03)	(4.04E-03)	(4.27E-03)

Continued ……

Table 2: Estimated Coefficients for the Generalized Almost Ideal Model With and Without Food Safety Variables (Continued ……)

	No Food Safety			With Food Safety		
	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}
Poultry Food Safety Variables						
k_{10}				-1.79E-03	-1.76E-03	-1.14E-03
				(1.20E-03)	(1.19E-03)	(1.13E-03)
k_{20}				-1.43E-03	-1.46E-03	-8.20E-04
				(1.25E-03)	(1.24E-03)	(1.19E-03)
k_{30}				-3.17E-03*	-3.37E-03*	-2.89E-03*
				(8.42E-04)	(8.52E-04)	(8.91E-04)
k_{11}				1.08E-02*	1.03E-02*	1.01E-02*
				(1.65E-03)	(1.64E-03)	(1.61E-03)
k_{21}				1.41E-02*	1.37E-02*	1.36E-02*
				(1.88E-03)	(1.86E-03)	(1.82E-03)
k_{31}				4.10E-03	4.09E-03	4.90E-03*
				(2.22E-03)	(2.20E-03)	(2.15E-03)
k_{12}				-7.22E-03*	-6.78E-03*	-7.05E-03*
				(1.77E-03)	(1.74E-03)	(1.72E-03)
k_{22}				-7.31E-03*	-6.84E-03*	-7.08E-03*
				(1.97E-03)	(2.00E-03)	(1.93E-03)
k_{32}				-2.45E-03	-2.22E-03	-3.16E-03*
				(1.56E-03)	(1.57E-03)	(1.56E-03)
k_{13}				4.18E-03*	4.15E-03*	4.43E-03*
				(1.02E-03)	(9.74E-04)	(9.24E-04)
k_{23}				8.45E-04	8.27E-04	7.72E-04
				(1.23E-03)	(1.22E-03)	(1.16E-03)
k_{33}				-7.10E-04	-9.00E-04	-5.30E-04
				(1.03E-03)	(1.04E-03)	(1.05E-03)
Autocorrelation Corrections						
r		0.442*			-0.086	
		(0.084)			(0.081)	
r_{11}			0.144			-0.171
			(0.187)			(0.147)
r_{12}			-0.135			0.017
			(0.197)			(0.153)
r_{21}			0.159			0.092
			(0.163)			(0.162)
r_{22}			0.629*			0.147
			(0.180)			(0.175)

Notes: Number in parentheses are the estimated standard errors and a * denotes coefficients that are statistically significantly different from zero at the 5% level.

Table 3: Estimated Coefficients for the Rotterdam Model With and Without Food Safety Variables

	No Food Safety			With Food Safety		
	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}
LL	*520.560*	*524.574*	*525.249*	*534.872*	*539.243*	*541.340*
t_1	-0.035*	-0.034*	-0.034*	-0.032*	-0.033*	-0.034*
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
t_2	0.022*	0.022*	0.022*	0.020*	0.021*	0.021*
	(0.002)	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)
q_1	0.061*	0.060*	0.060*	0.055*	0.056*	0.057*
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.005)
q_2	0.042*	0.042*	0.042*	0.040*	0.041*	0.041*
	(0.003)	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)
q_3	0.030*	0.030*	0.030*	0.028*	0.029*	0.030*
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
q_4	-0.033*	-0.032*	-0.032*	-0.028*	-0.029*	-0.029*
	(0.003)	(0.004)	(0.004)	(0.003)	(0.004)	(0.004)
q_5	-0.036*	-0.036*	-0.036*	-0.034*	-0.034*	-0.035*
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
q_6	-0.019*	-0.019*	-0.019*	-0.016*	-0.017*	-0.018*
	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)	(0.003)
g_{11}	-0.126*	-0.133*	-0.136*	-0.105*	-0.117*	-0.119*
	(0.034)	(0.030)	(0.030)	(0.035)	(0.030)	(0.028)
g_{12}	0.117*	0.118*	0.118*	0.106*	0.110*	0.114*
	(0.026)	(0.023)	(0.022)	(0.026)	(0.022)	(0.020)
g_{22}	-0.130*	-0.128*	-0.125*	-0.123*	-0.123*	-0.126*
	(0.025)	(0.021)	(0.021)	(0.026)	(0.022)	(0.020)
b_1	0.597*	0.589*	0.582*	0.547*	0.543*	0.554*
	(0.062)	(0.060)	(0.060)	(0.059)	(0.059)	(0.058)
b_2	0.264*	0.275*	0.278*	0.315*	0.309*	0.320*
	(0.051)	(0.050)	(0.050)	(0.047)	(0.047)	(0.046)
Beef Food Safety Variables						
f_{10}				3.29E-03	1.97E-03	1.18E-03
				(1.87E-03)	(1.83E-03)	(1.94E-03)
f_{11}				-1.90E-04	-3.80E-04	-1.20E-04
				(1.84E-03)	(1.73E-03)	(1.75E-03)
f_{12}				-1.67E-03	-2.11E-03	-2.78E-03
				(1.80E-03)	(1.68E-03)	(1.72E-03)
f_{13}				-2.93E-03	-1.98E-03	-8.20E-04
				(1.74E-03)	(1.64E-03)	(1.60E-03)
f_{20}				-2.48E-03	-2.14E-03	-8.60E-04
				(1.54E-03)	(1.44E-03)	(1.54E-03)
f_{21}				-4.20E-04	-1.30E-04	-3.80E-04
				(1.50E-03)	(1.40E-03)	(1.44E-03)
f_{22}				7.76E-04	8.17E-04	1.57E-03
				(1.50E-03)	(1.37E-03)	(1.41E-03)
f_{23}				3.14E-03*	2.16E-03	1.70E-03
				(1.42E-03)	(1.34E-03)	(1.33E-03)

Continued ……

Table 3: Estimated Coefficients for the Rotterdam Model With and Without Food Safety Variables (Continued..)

	No Food Safety			With Food Safety		
	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}
Pork Food Safety Variables						
p_{10}				9.85E-04	4.32E-04	-3.08E-06
				(1.76E-03)	(1.60E-03)	(1.51E-03)
p_{11}				1.26E-03	1.59E-03	2.36E-03
				(1.34E-03)	(1.30E-03)	(1.30E-03)
p_{12}				9.08E-04	1.00E-03	1.07E-03
				(1.11E-03)	(8.98E-04)	(8.85E-04)
p_{13}				-5.40E-04	-1.10E-04	5.89E-04
				(8.47E-04)	(8.16E-04)	(8.07E-04)
p_{20}				-3.30E-04	3.48E-04	6.49E-04
				(1.43E-03)	(1.29E-03)	(1.20E-03)
p_{21}				-1.02E-03	-1.65E-03	-2.08E-03
				(1.08E-03)	(1.04E-03)	(1.07E-03)
p_{22}				-5.60E-04	-6.50E-04	-6.10E-04
				(8.88E-04)	(7.15E-04)	(7.04E-04)
p_{23}				6.99E-04	7.40E-05	-3.00E-04
				(6.79E-04)	(6.49E-04)	(6.71E-04)
Poultry Food Safety Variables						
k_{10}				-5.37E-03	-3.92E-03	-3.01E-03
				(3.28E-03)	(3.08E-03)	(3.07E-03)f
k_{11}				-2.00E-03	-2.14E-03	-2.88E-03
				(3.26E-03)	(3.08E-03)	(3.13E-03)
k_{12}				2.37E-03	2.78E-03	3.44E-03
				(3.28E-03)	(3.09E-03)	(3.13E-03)
k_{13}				3.08E-03	2.16E-03	3.75E-04
				(3.24E-03)	(3.07E-03)	(2.97E-03)
k_{20}				2.50E-03	1.68E-03	7.24E-04
				(2.65E-03)	(2.47E-03)	(2.43E-03)
k_{21}				3.05E-03	3.50E-03	4.01E-03
				(2.65E-03)	(2.48E-03)	(2.52E-03)
k_{22}				-2.70E-03	-2.97E-03	-3.72E-03
				(2.67E-03)	(2.49E-03)	(2.53E-03)
k_{23}				-5.25E-03*	-3.89E-03	-2.79E-03
				(0.00262)	(0.00246)	(0.00245)
Autocorrelation Corrections						
r		-0.247*			-0.299*	
		(0.089)			(0.092)	
r_{11}			-0.227			-0.449*
			(0.213)			(0.220)
r_{12}			0.144			-0.023
			(0.257)			(0.248)
r_{21}			0.068			0.295
			(0.174)			(0.170)
r_{22}			-0.257			-0.151
			(0.213)			(0.193)

Notes: Number in parentheses are the estimated standard errors and a * denotes coefficients that are statistically significantly different from zero at the 5% level.

Table 4: Hypothesis Tests for the Significance of Food Safety Variables and Autocorrelation Corrections

Table 4a: Testing Alternative Lag Lengths for Food Safety Variables in the Presence of Alternative Autocorrelation Corrections Using the GAI model

Autocorrelation Correction	Hypothesis Test (H_0 vs. H_a) m=0 v No-FS	m=1 v m=0	m=2 v m=1	m=3 v m=2	m=4 v m=3	m=3 v No-FS
N-Rmatrix	38.915*	19.125*	11.932	31.035*	13.165	88.043*
D-Rmatrix	24.407*	18.449*	12.249	30.579*	12.716	77.446*
F-Rmatrix	20.432*	18.230*	13.609	30.240*	12.546	74.652*
df	9	9	9	9	9	36
$\chi_{0.05,df}$	16.919	16.919	16.919	16.919	16.919	43.773

Table 4b: Testing Alternative Autocorrelation Corrections in the Presence of Alternative Lag Lengths for Food Safety Variables Using the GAI model

Hypothesis Test (H_0 vs. H_a)	df	$\chi_{0.05,df}$	Alternative Lag Lengths of Food Safety Variables No-FS	m=0	m=1	m=2	m=3	m=4
N-R^{matrix} vs. D-R^{matrix}	1	3.841	15.984*	0.533	0.019	0.479	0.396	0.102
N-R^{matrix} vs. F-R^{matrix}	4	9.488	19.647*	1.036	0.886	3.052	3.606	3.339
D-R^{matrix} vs. F-R^{matrix}	3	7.815	4.107	0.519	0.867	2.591	3.227	3.242

Table 4c: Testing Alternative Lag Lengths for Food Safety Variables in the Presence of Alternative Autocorrelation Corrections Using the Rotterdam model

Autocorrelation Correction	Hypothesis Test (H_0 vs. H_a) =0 v No-FS	m=1 v m=0	m=2 v m=1	m=3 v m=2	m=4 v m=3	m=3 v No-FS
N-Rmatrix	7.890	3.730	5.814	5.186	11.700	20.600
D-Rmatrix	8.051	4.615	6.784	3.665	11.966	20.893
F-Rmatrix	8.806	5.224	7.497	3.143	9.431	22.144
df	6	6	6	6	6	24
$\chi_{0.05,df}$	12.592	12.592	12.592	12.592	12.592	36.415

Table 4d: Testing Alternative Autocorrelation Corrections in the Presence of Alternative Lag Lengths for Food Safety Variables Using the Rotterdam model

Hypothesis Test (H_0 vs. H_a)	df	$\chi_{0.05,df}$	Alternative Lag Lengths of Food Safety Variables No-FS	m=0	m=1	m=2	m=3	m=4
N-R^{matrix} vs. D-R^{matrix}	1	3.841	7.177*	7.042*	7.585*	8.183*	6.226*	6.227*
N-R^{matrix} vs. F-R^{matrix}	4	9.488	8.170	8.939	10.080*	11.406*	8.876	6.549
D-R^{matrix} vs. F-R^{matrix}	3	7.815	1.176	2.085	2.710	3.468	2.849	0.534

Notes: m denotes the lag length of food safety variables included in each model; No-FS denotes a model with no food safety variables included; and df denotes degrees of freedom. All likelihood ratio test statistics are calculated using the adjusted likelihood ratio test statistic: $LR^A = [(T-k)/T]*2*(LL^U-LL^R)$ where T is the sample size, k is the estimated number of parameters in the unrestricted model; LL^U and LL^R are the maximized likelihood value in the unrestricted and restricted models, respectively. A * denotes a significant test statistic at the 5% level.

Table 5: Estimated Price, Expenditure and Food Safety Elasticities With Alternative Autocorrelation Corrections

	GAI Model			Rotterdam Model		
	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}
Price						
h_{11}	-0.886	-0.885	-0.893	-0.754	-0.774	-0.789
h_{12}	-0.031	-0.025	-0.027	-0.105	-0.094	-0.094
h_{13}	-0.074	-0.067	-0.064	-0.215	-0.199	-0.206
h_{21}	-0.264	-0.267	-0.246	-0.194	-0.169	-0.176
h_{22}	-0.824	-0.827	-0.817	-0.741	-0.736	-0.758
h_{23}	-0.368	-0.376	-0.367	-0.154	-0.165	-0.174
h_{31}	0.235	0.242	0.234	-0.369	-0.354	-0.303
h_{32}	-0.194	-0.204	-0.215	-0.115	-0.151	-0.120
h_{33}	-0.275	-0.282	-0.301	-0.226	-0.253	-0.221
Expenditure						
h_{1M}	0.990	0.977	0.984	1.074	1.066	1.089
h_{2M}	1.457	1.469	1.430	1.089	1.070	1.108
h_{3M}	0.234	0.244	0.283	0.710	0.758	0.643
Food Safety						
Immediate Effect						
w_{11}	0.01013	0.01012	0.01005	0.00646	0.00387	0.00232
w_{21}	-0.01553	-0.01575	-0.01496	-0.00856	-0.00738	-0.00297
w_{31}	-0.00195	-0.00163	-0.00258	-0.00419	0.00085	-0.00165
w_{12}	-0.00879	-0.00857	-0.01020	0.00193	0.00085	-0.00001
w_{22}	0.00448	0.00392	0.00598	-0.00115	0.00120	0.00224
w_{32}	0.01372	0.01404	0.01482	-0.00335	-0.00401	-0.00332
w_{13}	-0.00250	-0.00230	-0.00103	-0.01054	-0.00770	-0.00591
w_{23}	0.01337	0.01365	0.01122	0.00865	0.00581	0.00250
w_{33}	-0.01368	-0.01455	-0.01383	0.01476	0.01152	0.01176
Total Effect						
m_{11}	0.00265	0.00243	0.00151	-0.00293	-0.00491	-0.00499
m_{21}	-0.00915	-0.00948	-0.00785	0.00354	0.00244	0.00702
m_{31}	0.00558	0.00652	0.00632	0.00241	0.00924	0.00264
m_{12}	-0.02406	-0.02469	-0.02454	0.00513	0.00571	0.00789
m_{22}	0.02939	0.02948	0.02890	-0.00421	-0.00650	-0.00811
m_{32}	0.01350	0.01490	0.01528	-0.00717	-0.00529	-0.00860
m_{13}	0.02155	0.02230	0.02370	-0.00376	-0.00220	-0.00408
m_{23}	-0.00170	-0.00167	-0.00417	-0.00829	-0.00578	-0.00614
m_{33}	-0.04504	-0.04683	-0.04638	0.02220	0.01437	0.01982

Continued ......

Table 5: Estimated Price, Expenditure and Food Safety Elasticities With Alternative Autocorrelation Corrections (Continued ….)

	GAI Model			Rotterdam Model		
	N–R^{matrix}	D–R^{matrix}	F–R^{matrix}	N–R^{matrix}	D–R ^{matrix}	F–R^{matrix}
Percent of Observations Consistent with Prior Beliefs						
Ph_{11}	96.970	96.970	96.970	100.000	100.000	100.000
Ph_{22}	100.000	100.000	100.000	100.000	100.000	100.000
Ph_{33}	95.455	95.455	95.455	100.000	100.000	100.000
Pw_{11}	1.515	1.515	3.030	0.000	0.000	0.000
Pw_{22}	1.515	0.000	1.515	100.000	0.000	0.000
Pw_{33}	93.939	92.424	96.970	0.000	0.000	0.000
Pm_{11}	34.849	34.849	34.849	100.000	100.000	100.000
Pm_{22}	18.182	18.182	18.182	100.000	100.000	100.000
Pm_{33}	83.333	83.333	78.788	0.000	0.000	0.000
P_{NSD}	93.939	93.939	92.424	100.000	100.000	100.000

Notes: h_{ij} represent the Marshallian price elasticities of demand for the i^{th} good with respect to the j^{th} price, and h_{iM} is expenditure elasticities for the i^{th} good, where i=1 for beef, 2 for pork, and 3 for poultry. Ph_{ii} , Pw_{ii} , and Pm_{ii} and measure the percentages of observations that satisfy h_{ii}< 0 , w_{ii} < 0, and m_{ii} < 0. P_{NSD} is the percentage of observations that satisfy the curvature requirements of negative semi-definiteness of the Slutsky matrix. Figures shown are the sample means of the elasticities computed at every data point using predicted expenditure shares.

Figure 1: Food Safety Variables

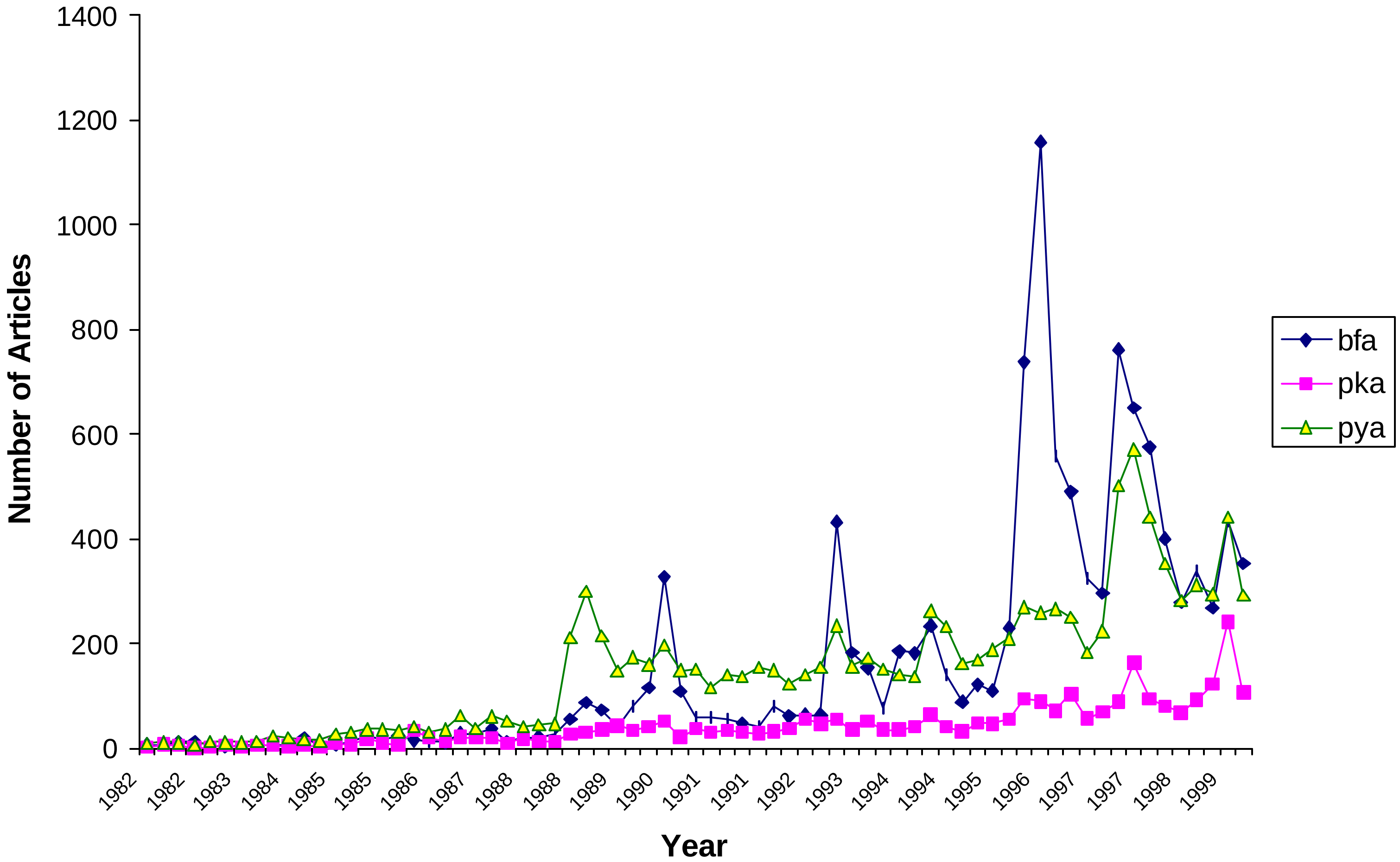


Figure 2: Estimates of m_{11} Across Alternative Models

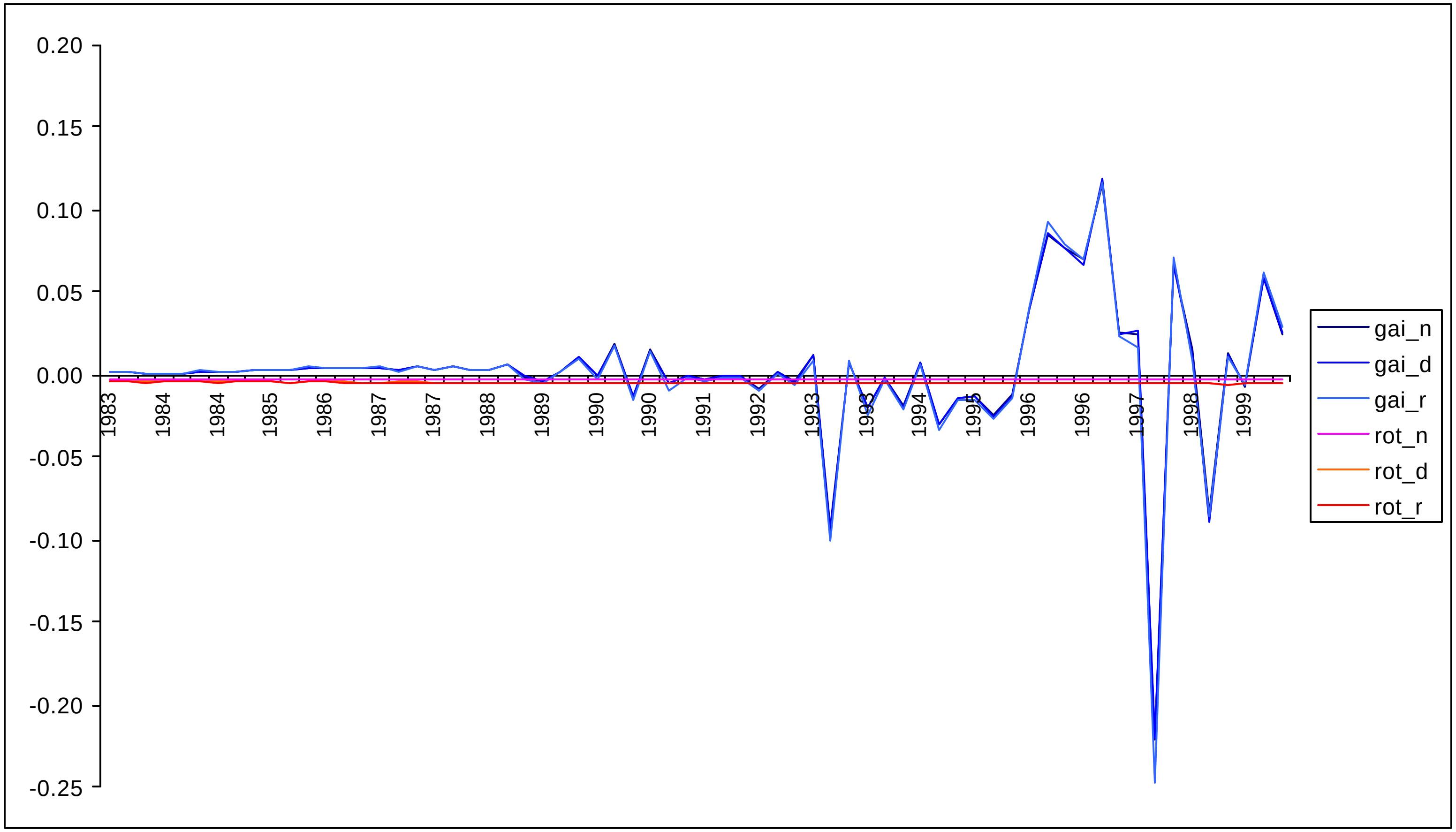


Figure 3: Estimates of m_{22} Across Alternative Models

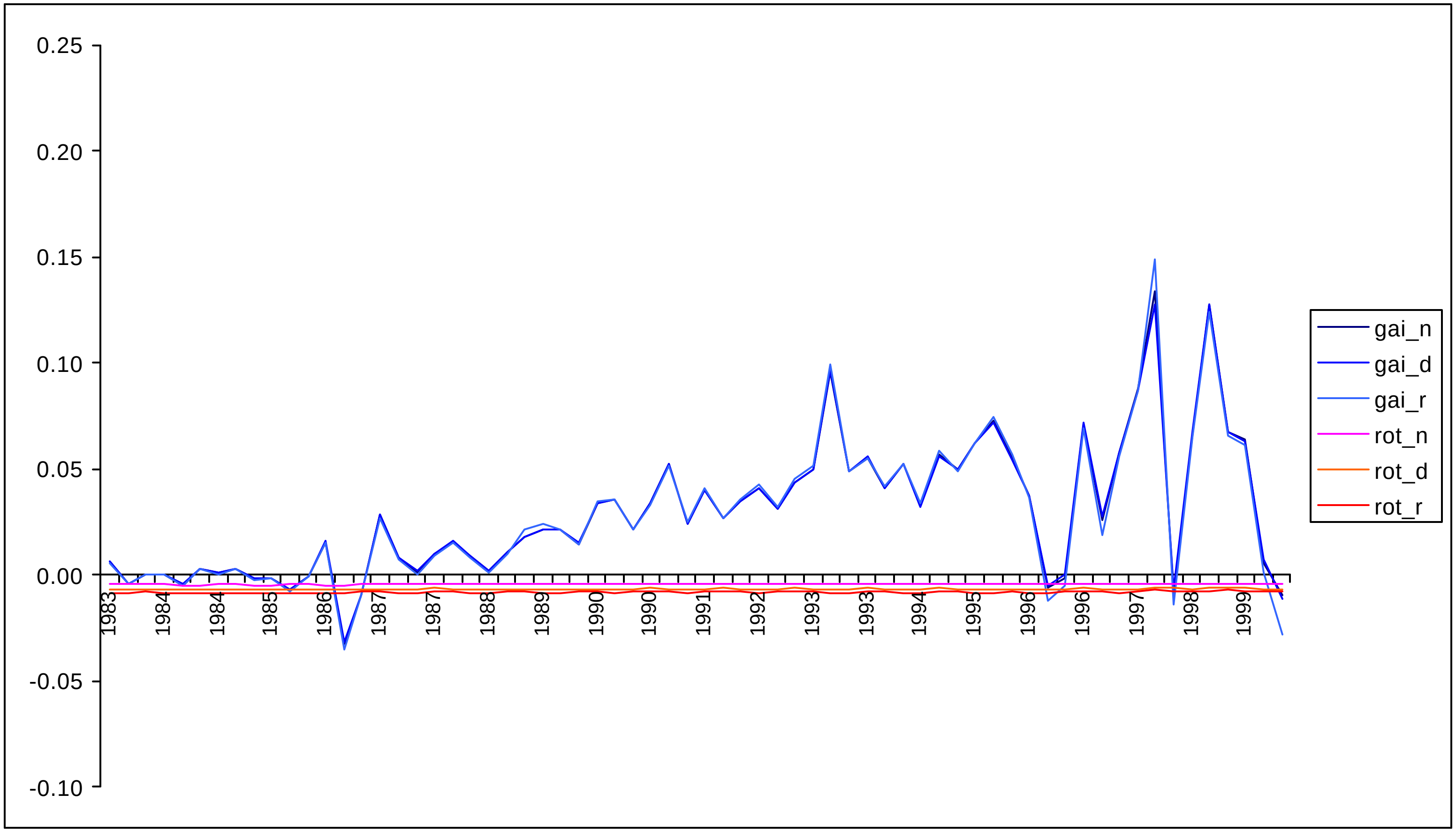
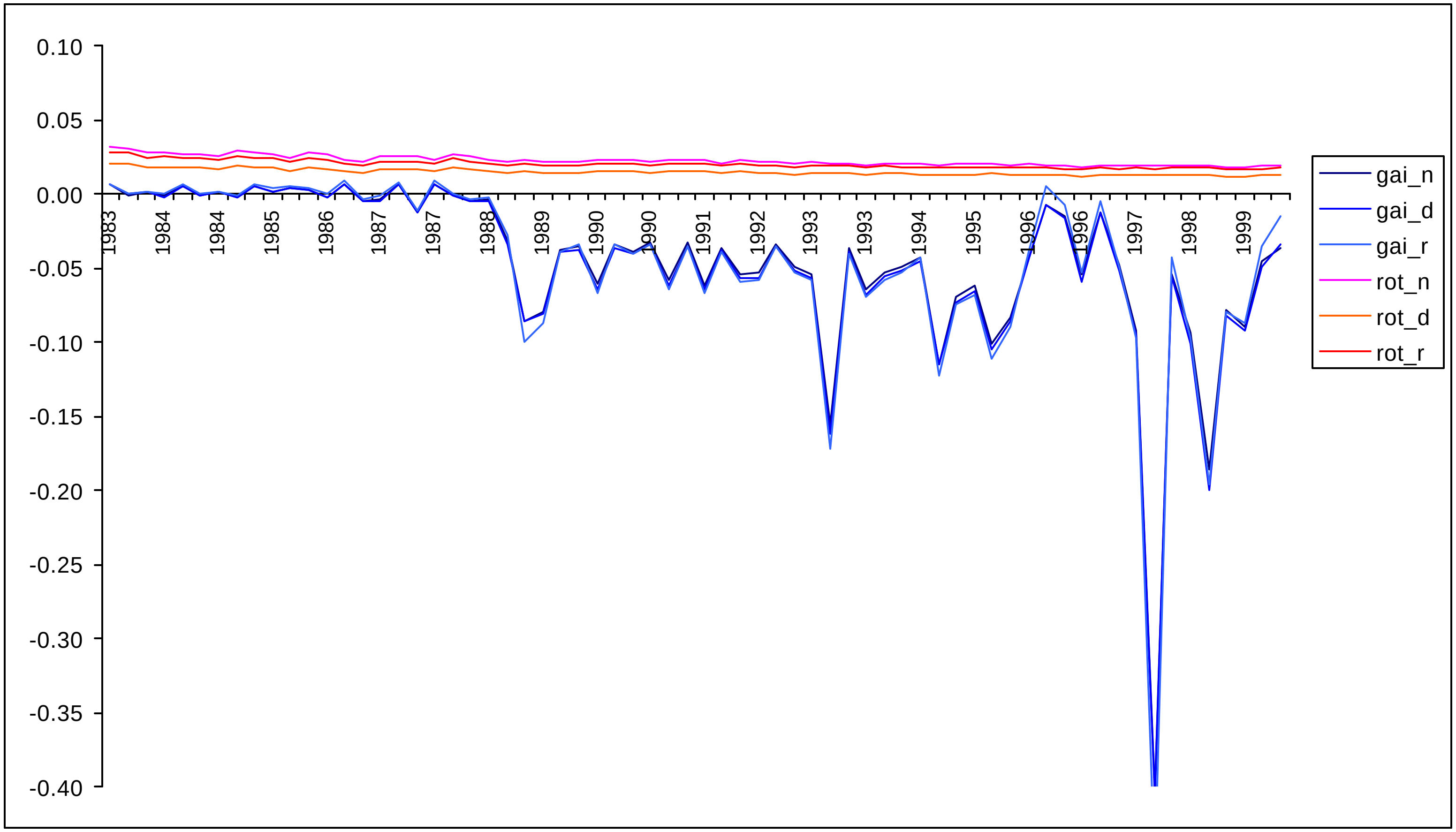


Figure 4: Estimates of m_{33} Across Alternative Models



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