A Test for the Consistency of Demand Data with Consumer Preference Theory

Yong Sakong and Dermot J. Hayes

Working Paper 92-WP 87
January 1992

Center for Agricultural and Rural Development
Iowa State University
Ames, Iowa 50011

Yong Sakong is a postdoctoral research associate, Department of Economics; and Dermot J. Hayes is an associate professor of economics and head of the Trade and Agricultural Policy Division, CARD.

We are indebted to Tae-Kyun Kim for his helpful suggestions.

ABSTRACT

We introduce adding-up, convexity, and noninferiority into the weak axiom of revealed preferences and show how we can use this approach to detect and measure structural change without estimation or subjective input. We use this new approach to measure taste changes among meats for the United States, Canada, Japan, and South Korea.

Introduction

Recent papers by Alston and Chalfant (1991a, 1991b, 1991c), Chalfant and Alston (1988), and Cox and Chavas (1987, 1990) have cast doubt on the parametric techniques previously used for measuring taste changes. Alston and Chalfant (1991b) most graphically demonstrated the problem with parametric methods by showing that one could detect structural change in beef demand almost 100 percent of the time in a system in which, by construction, no such change exists. In their paper, data were generated by a linear demand system, and the structural change test was conducted with a logarithmic and Almost Ideal Demand System (AIDS) model.

This issue is very similar to that described by Leamer (1983). Taste change effects, if they exist, will be relatively small. To measure these changes by parametric methods, the researcher is forced to make decisions about the functional form and the estimation procedure. If the researcher is looking for (or if the system rewards) evidence of taste change, he or she need only search among the set of inferences one can draw from a particular data set for those results that are most pleasing.

The proposed solution in the nonparametric literature is to avoid to the greatest possible extent all decisions that might influence the possible outcome. For demand analysis, this solution simply involves testing the data for consistency with the weak or strong axioms of revealed preference. These axioms avoid the need to express and estimate the direct, indirect, or expenditure functions and instead rely on very intuitive conditions. These conditions state that, in the absence of taste change, if a bundle of Q_1 is revealed preferred to a bundle of Q_2 at one point, then Q_2 cannot be revealed preferred to Q_1 at another point, unless taste change has occurred.[1]

One problem that occurs with real data is that, as real income has increased, real expenditures on some commodity bundles have also increased. This trend makes it very difficult to find a bundle that was affordable and not consumed in one period but that was consumed later although the previous

[1]Consider a batch of commodities, Q_1, purchased at prices P_1. Now consider a second commodity bundle, Q_2, such that $P_1Q_2 \leq P_1Q_1$. Because the consumer could afford Q_2 at prices P_1 but chose Q_1 instead, Q_1 is revealed preferred to Q_2. The weak axiom states that, at prices P_2, we will not see $P_2Q_1 \leq P_2Q_2$, i.e., Q_2 will not be revealed preferred to Q_1 at any set of prices. The strong axiom introduces transitivity, i.e., if Q_1 is revealed preferred to Q_2 and Q_2 is revealed preferred to Q_3, then Q_3 must never be revealed preferred to Q_1.

bundle remained affordable. This can best be demonstrated in Figure 1, where the left panel shows a violation of consumer preference theory and the right panel shows a similar situation but with the budget constraint shifted out.

Intuitively, tastes have shifted away from good q_1 to good q_2 in both situations. For a violation of consumer preference, however, the budget lines must cross. Hence, the situation in the right panel does not provide evidence of a violation of consumer preference theory.[2] The movement toward q_2 could be explained by an almost vertical Engel curve, and the movement away from q_1 might have occurred because q_2 is an inferior good. This situation motivates the study that follows.

Consider Figure 2. Here, as discussed earlier, the movement from Q_1 to Q_2 can be explained by assuming that q_2 is an inferior good. Suppose that we are prepared to assume that q_2 is not an inferior good, then the question arises whether the consumption change is consistent with consumer preference theory. If consumption patterns are not consistent, then we may conclude that tastes (or consumer preferences) have changed. One very obvious implication of consistent preferences is that if consumers are held on the same indifference curve and subjected to the same price vector, then they should consume the same bundle. However, we can show that so long as q_2 is not inferior, the range of possible bundles found by adjusting for the price effect from Q_1 does not overlap with the range found when the expenditure effect is compensated from Q_2.

To show that two regions cannot overlap, consider how actual demand would change if we start at Q_1 and change the prices to P_2 and then compensate the consumer for the price change. Because we do not know the shape of the indifference curve, we draw the new budget line $A'B'$ to allow for the maximum possible compensation, i.e., we allow the consumer to purchase Q_1 at the new price line. In reality, the new budget line will lie to the left of $A'B'$ and the new compensated bundle, $h(U_1, P_2)$, will lie in the region AQ_1A'. Now if we start at Q_2 and compensate for the price increase, the new bundle, $g(U_1, P_2)$, will lie in the region DQ_2E. This is true because the consumer

[2]Alston and Chalfant (1991a) recently showed that the probabilities of violating WARP tend to increase as the size of the taste change increases and as the growth rate of total expenditures decreases.

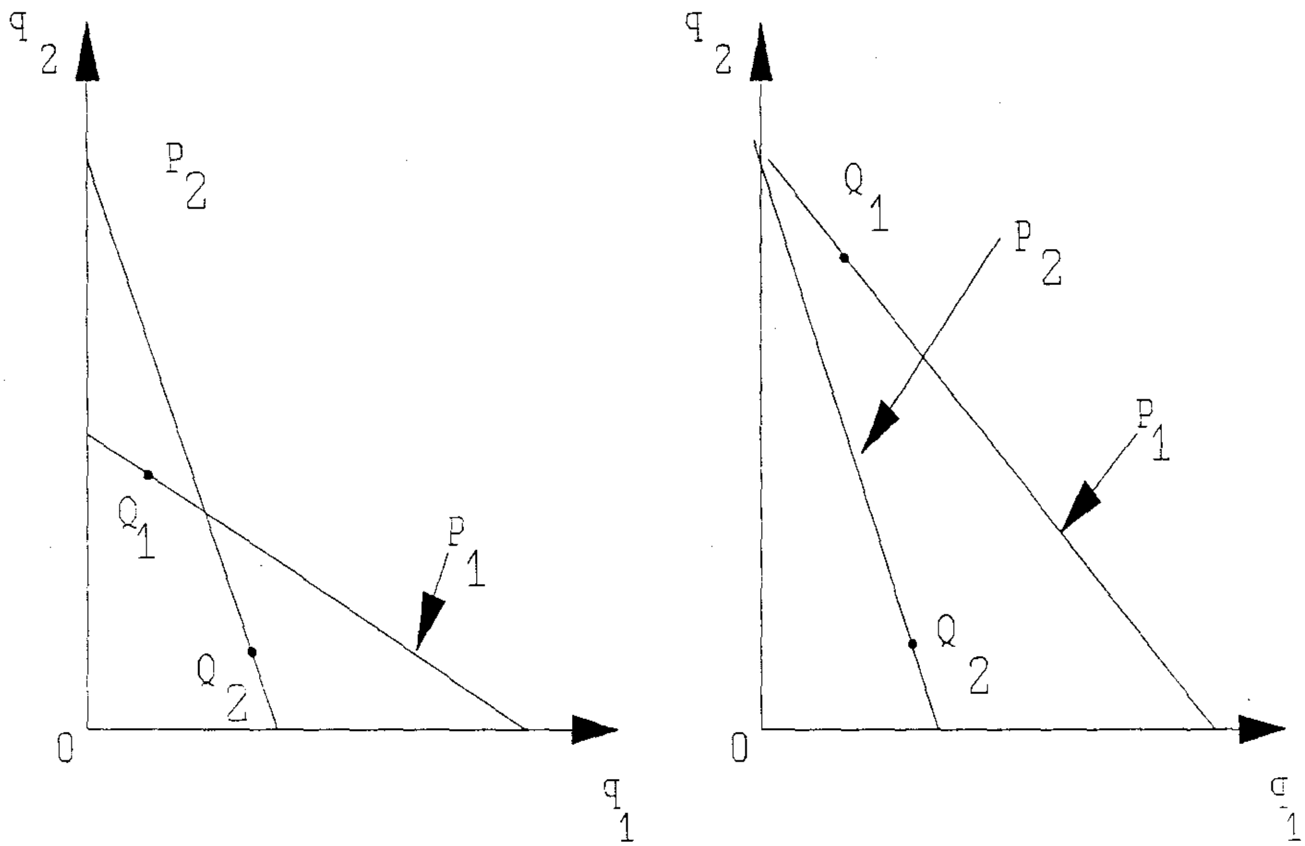

Figure 1. A two-good example of WARP

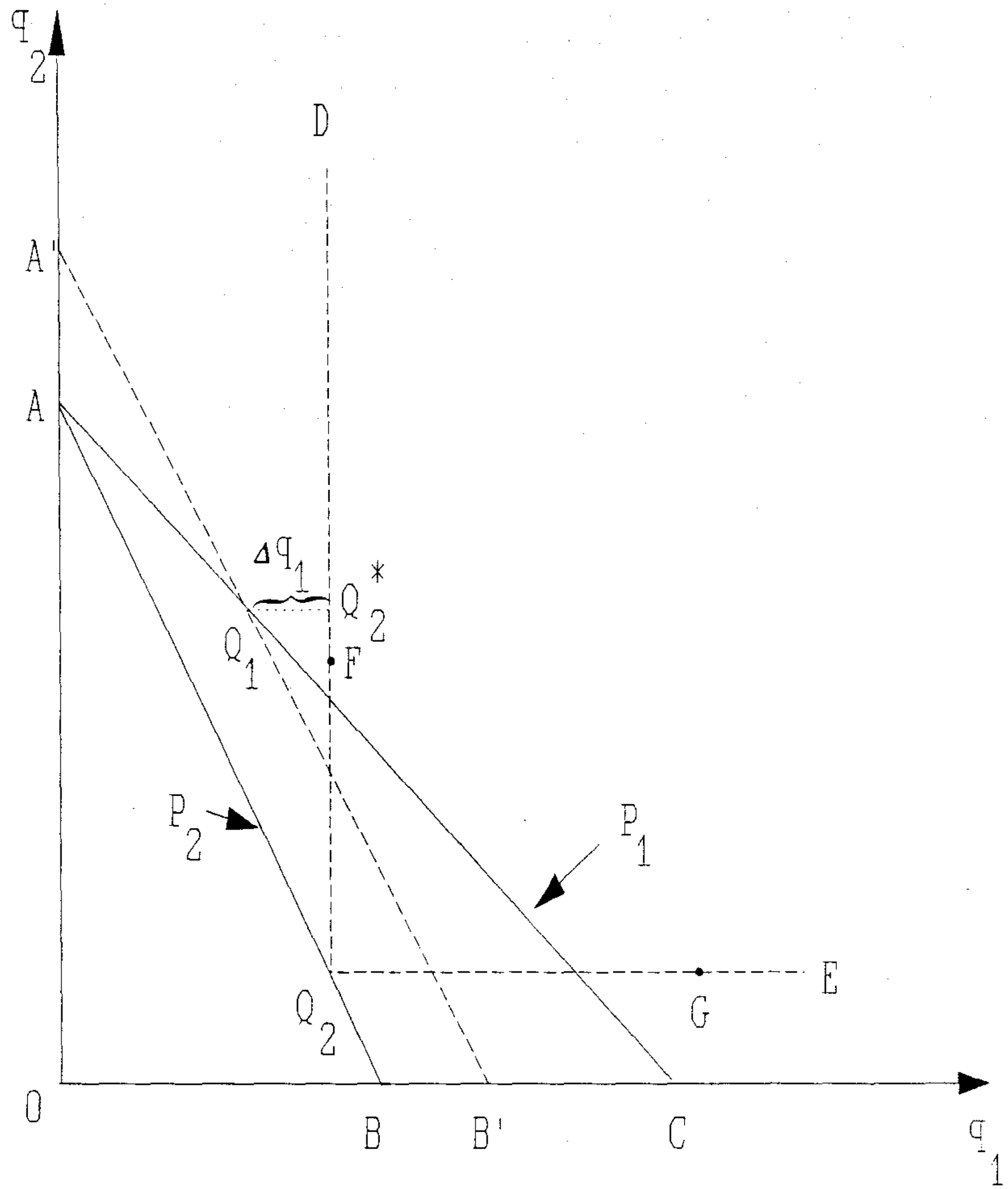

Figure 2. An example of a violation of consumer perferences

will spend the additional compensatory income on both non-inferior goods. We have created asituation with two demands, $h(U_1, P_2)$ and $g(U_1, P_2)$. If consumer preferences are consistent, then the two solutions should be identical. Yet as we have shown graphically, the two regions do not overlap.

To measure this change in tastes, we ask which set of expenditure elasticities best explains the behavior, the remainder being attributed to taste change. This gives us the minimum taste change that explains the data. In this very simple example, one might conclude that the income elasticities are such that all additional compensatory income is spent on q_2; this would lead to Q_2^*. The minimum taste change is therefore away from q_1 by the horizontal amount $Q_1 - Q_2^*$ measured in units of q_1 as shown by Δq_1 in Figure 2. This example motivates the test that follows. We minimize the degree of taste change that satisfies both consistency of preferences and the restrictions we place on these expenditure elasticities. This procedure is much more complicated than that indicated by Figure 2 because we have ignored the adding-up condition. However, the intuition remains the same.

In the analysis that follows we progressively impose restrictions on the slopes of Engel curves for meat data from the United States, Canada, Japan, and South Korea. First we impose non-negativity, adding-up, and convexity. Then, reasonable ranges for the expenditure elasticities are imposed. Finally, we impose restrictions on how the expenditure elasticities can change from year to year.[3] In all cases, we simultaneously estimate the minimum consumption changes needed to satisfy consistency of preferences and the expenditure elasticities that best explain this behavior. The mechanism used to measure the degree to which preference consistency is violated is based on a linear programming model recently developed by Cox and Chavas (1987, 1990). We introduce to their model the modifications needed to simultaneously impose restrictions on expenditure elasticities, the adding-up condition, and indifference curve convexity. In the empirical analysis, we use the new method, which we call a test for consistency of preferences, to detect and measure taste change in the

---

[3] We can test a taste change by imposing all the restrictions including WARP, adding-up, non-negativity, reasonable range of expenditure elasticities, and year-to-year changes of expenditure elasticities.

United States, Canada, Japan, and South Korea.

A Test for Consistency of Preferences

Suppose that n goods exist and that demand for good i, q_i, is a function of prices, income, and taste changes:

$$q_i = q_i(P', x, TC') \tag{1}$$

where P and TC are a price vector and a taste change vector of n goods, respectively; i.e., $P' = (p_1, p_2, \cdots, p_n)$ and $TC' = (tc_1, tc_2, \cdots, tc_n)$, and x is income (or expenditure). If we differentiate equation (1), we find:

$$\Delta q_i = \sum_{j=1}^{n} \frac{\partial q_i}{\partial p_j}\Delta p_j + \frac{\partial q_i}{\partial x}\Delta x + \sum_{j=1}^{n} \frac{\partial q_i}{\partial tc_j}\Delta tc_j \,. \tag{2}$$

Using the Slutsky equation and temporarily assuming that $\Delta tc_j = 0$ for all $j = 1, 2, \cdots, n$, equation (2) can be rewritten as:

$$\Delta q_i = \sum_{j=1}^{n} \frac{\partial q_i}{\partial p_j}\Big|_{U_0}\Delta p_j + a\frac{q_i}{x}\epsilon_{ix} \tag{3}$$

where $a \equiv \Delta x - \sum_{j=1}^{n} q_j \Delta p_j$ and ϵ_{ix} is an expenditure elasticity of good i.

Equation (3) separates the demand change induced by price changes and expenditure change into two effects: the first part is the substitution effect induced by price changes and the second is the expenditure effect induced by both price changes and expenditure changes. Subtracting the second part in the right side of equation (3) from the observed demand data, $q_i^* = q_i - a(q_i/x)\epsilon_{ix}$, reflects only the substitution effect. By holding the consumer on the same indifference curve in this manner, we can respecify the conditions under which consistency of preferences is violated as $P_t'Q_t^* > P_t'Q_s^*$ for all t and s, where Q_t^* is a compensated consumption bundle at time t, i.e., $Q_t^* = (q_{1t}^*, \ldots, q_{nt}^*)$.[4] To see why this is true, consider Figure 3, with two consumption bundles, Q_1 and Q_2^*. Bundle Q_1 is a base-year consumption bundle and Q_2^* is the optimal consumption bundle at time 2

---

[4] If t or s is the base year, Q^* is a consumption bundle rather than a compensated consumption bundle.

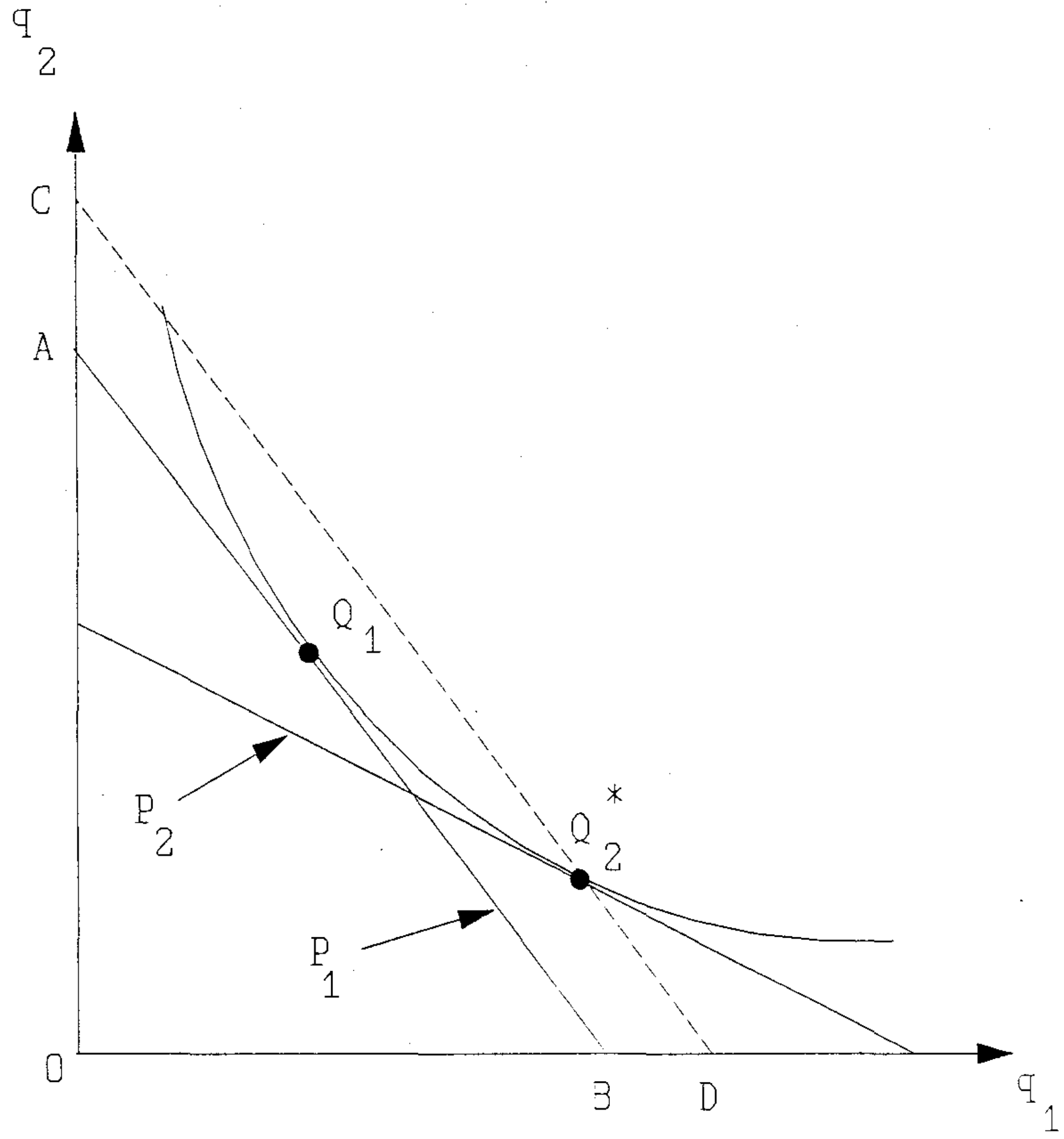

Figure 3. Graphical representation of the convexity condition

prices and time 1 utility level. Note that if the indifference curve is convex, then all Q_i^* will lie to the right of AB. Now if we express Q_2^* at time 1 prices, then expenditure $P_1Q_2^*$ will be equal to or greater than P_1Q_1. To see why this is true, draw a line through Q_2^* parallel to AB (CD in Figure 3) and measure expenditure in terms of good 2. If $P_1Q_2^*$ is less than P_1Q_1, then preferences are inconsistent with the data. Because the inequality $P_1{}'Q_1 \leq P_1{}'Q_2^*$ depends on the convexity of the indifference curve, we call this the convexity condition.[5]

By using the consistency WARP test, some values of TC always exist to satisfy the inequality $P_t'(Q_t^* - TC_t) \leq P_t(Q_s^* - TC_s)$ for all t and s.[6] We can find the minimum TC that satisfies these inequalities by solving the following problem:[7]

$$\min_{TC,\phi} \; b'TC$$

$$\text{s.t. } P_t\,(Q_t^* - TC_t) \leq P_t\,(Q_s^* - TC_s) \quad \textit{for all } t \textit{ and } s,$$

$$\sum_i w_i^t \epsilon_{ix}^t = 1 \quad \textit{for all } t, \textit{ and} \tag{4}$$

$$\epsilon_{ix}^t \geq 0 \quad \textit{for all } i \textit{ and } t$$

where ϕ is a vector of expenditure elasticities, b is arbitrarily defined such that problem (4) is bounded, and w_{it} and $\epsilon_{ix}{}^t$ are an expenditure share and an expenditure elasticity of good i at time t. The first and second constraints represent the consistency and adding-up condition, respectively. The third constraint represents the non-negativity of expenditure elasticities.

To see how the adding-up condition influences the results, consider that the compensated demands of goods 1 and 2 at time 2, q_1^* and q_2^*, are:

---

[5] Varian (1984) shows the same condition for a cost minimizing firm.

[6] Taste change may be negative or positive so that $(TC^+ - TC^-)$ is actually substituted into TC in the linear programming problem, where $TC^+ \geq 0$ and $TC^- \geq 0$.

[7] Chavas and Cox (1990) test for technical changes by using a similar method.

$$q_1^* = q_1 - \frac{a}{x} q_1 \epsilon_{1x} \tag{5}$$

and

$$q_2^* = q_2 - \frac{a}{x} q_2 \epsilon_{2x} \tag{6}$$

and the adding-up condition is

$$w_1 \epsilon_{1x} + w_2 \epsilon_{2x} = 1 \ . \tag{7}$$

Because equations (5) through (7) have four unknown variables and three equations, we can obtain the following relationship:

$$q_2^* = \Pi_1 - \Pi_2 q_1^* \tag{8}$$

$$\textit{where,} \quad \Pi_1 = q_2 (1 - \frac{a}{x w_2} + \frac{w_1}{w_2})$$

$$\Pi_2 = \frac{w_1}{w_2} \frac{q_2}{q_1} \ .$$

Because Π_1 and Π_2 are known coefficients and Π_2 is positive, the compensated demand of good 2, q_2^*, has a linear negative relationship with q_1^*. Now suppose in Figure 2 that this relationship was satisfied along the line that connects Q_2^* and G. Then, the restriction has no impact as it allows the vertical move from Q_2 to Q_2^*. Suppose, however, that the expenditure elasticity that underlies the move from Q_2 to Q_2^* violated adding-up; this is equivalent to stating that the relationship between q_2^* and q_1^* passes below Q_2^*, for example, at point F. Then, the minimum taste change would be $Q_1 - Q_2^*$ units of q_1 and $Q_2^* - F$ units of q_2.

One practical problem remains: because Δ represents a very small change, Δx, Δq, and Δp in equations (2) through (8) must also be small changes. To address this, we set the year of first observations as a base year, or time 1. We then denote the partial expenditure effect of good i at time t (pie_{it}) as an expenditure effect that occurs when consumption bundles and their prices at time t are compared with those at a previous time ($t - 1$). Then the expenditure effect on good i at time t,

(ie_{it}), is the sum of partial expenditure effects from time 2 to time t:

$$ie_{it} = \sum_{j=2}^{t} pie_{ij} = \sum_{j=2} a_k \frac{q_{ij}}{x_j} \epsilon_{ix}^{j} \,. \tag{9}$$

Similarly, we let the partial taste change of good i at time t, ptc_{it}, be a taste change that occurred between $(t - 1)$ and t. Taste change of good i at time t is therefore also measured as the sum of the partial changes from time 2 to time t:

$$tc_{it} = \sum_{j=2}^{t} ptc_{ij} \,. \tag{10}$$

In effect, the taste change at time t would be the accumulation of past as well as current taste changes when the current consumption bundle is compared with that of the base year.

Suppose that there are two goods, A and B, and that a positive taste change occurs in one good and a negative taste change occurs in the other good. When two goods are assumed to be substitutes, then a positive taste change in one good will decrease demand for the other good. Unfortunately, we cannot distinguish whether a taste change in good A causes demand for good B to change or whether a taste change in good B causes demand for good A to change. Therefore, taste change in good i in our model can be explained by $\sum_{j=1}^{n} (\partial q_i/\partial tc_j)\Delta tc_j$ rather than only by $(\partial q_i/\partial tc_i)\Delta tc_i$. That is, the taste change of good i is measured by the changes in the demand for the good, which cannot be interpreted by its own substitution effect and expenditure effect, even though it may be caused by the taste changes in other goods. Therefore, the third term of the right side in equation (2) will be simply expressed as tc_i.

Substituting (8) and (9) into (2) and rearranging, the compensated demand for good i at time t is:

$$q_{it}^{*} = q_{it} - \sum_{j=2}^{t} a_j \frac{q_{ij}}{x_j} \epsilon_{ix}^{j} + \sum_{j=2}^{t} ptc_{ij} \,. \tag{11}$$

By substituting (11) into (4), we avoid the need for estimating q_{it}^{*}. The model we actually solve is:

$$\min_{PTC\phi} \; b' \, PTC \; ; \tag{12}$$

$$\begin{aligned}
s.t. \quad (i) \; & \sum_{i=1}^{n} p_{it} q_{it} - \sum_{i=1}^{n} p_{it} q_{is} \\
& \le \sum_{i=1}^{n} \sum_{j=2}^{t} a_j \frac{p_{it} q_{ij}}{x_j} \, \epsilon_{ix}^{j} - \sum_{i=1}^{n} \sum_{j=2}^{s} a_j \frac{p_{it} q_{ij}}{x_j} \, \epsilon_{ix}^{j} \\
& + \sum_{i=1}^{n} \sum_{j=2}^{t} p_{it} ptc_{ij} - \sum_{i=1}^{n} \sum_{j=2}^{s} p_{it} ptc_{ij} \\
& \textit{for all } t \textit{ and } s \; ; \\
(ii) \; & \sum_{i} w_i^t \epsilon_{ix}^t = 1 \quad \textit{for all } t \; ; \textit{ and} \\
(iii) \; & \epsilon_{ix}^t \ge 0 \quad \textit{for all } i \textit{ and } t \; .
\end{aligned}$$

Note that, for this model, we do not need data on the unobservable Q_1^*.

Empirical Application

Data

Data on per capita annual consumption of beef, pork, and chicken for 1971 through 1984 for four countries (the United States, Canada, Japan, and South Korea) are used. The U.S. data are taken from Chalfant, the Canadian data from Van Kooten (1987), and the Japanese data from Wahl and Hayes (1990). The South Korean data are collected from the annual reports of the Agricultural Cooperative Federation and the National Livestock Cooperatives Federation. Because of the enormous number of restrictions necessary to solve the model, we were limited to 15 years of data. We chose the 15 years that all four data sets had in common. This centers the U.S. data around the 1976 to 1978 period and therefore includes the years of maximum U.S. beef consumption as well as the decrease in consumption that triggered the series of taste change studies mentioned earlier.

The results obtained from U.S. meat demand data using the programming problem (12) are presented in Figure 4. This figure represents the per capita change in pounds from the base year of 1971. One of the more interesting features of these results is the gradual trend away from beef. As

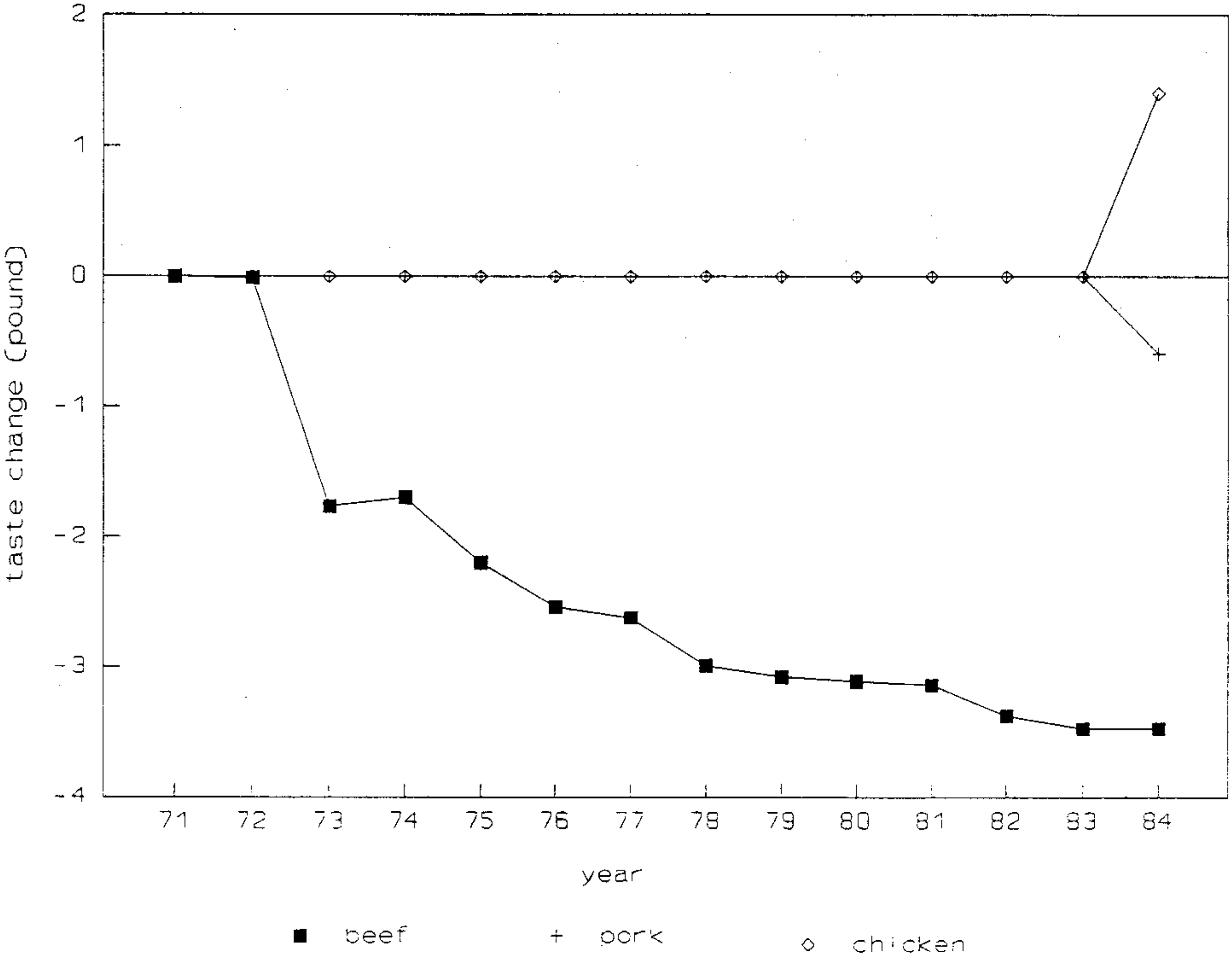

Figure 4. Taste change in U.S. meat demand

the program is written, each year is treated independently; therefore, years in which taste moved in favor of beef can in practice be followed by years in which the movement was against beef. The existence of a trend away from beef would seem to indicate that the source of the inconsistency—be it data driven, or caused by health concerns—is not random.

The beef results indicate a cumulative movement away from beef of approximately 3.5 pounds per capita with most (2 pounds) occurring from 1972 to 1973. Actual per capita U.S. beef consumption from 1971 to 1980 was 83.4, 85.4, 80.5, 85.6, 87.9, 94.4, 91.8, 87.2, 78.0 and 76.5 pounds, respectively. The results for U.S. pork and chicken consumption are consistent with theory, with small violations against pork and towards chicken in 1984.

In this model we imposed only adding-up and non-negative expenditure elasticities. The expenditure elasticities that underlie Figure 4 are shown in Table 1. As mentioned, these elasticities are found by minimizing the amount of taste change. The program makes no attempt to realistically measure these elasticities other than to ensure that they satisfy adding-up and non-negativity. The expenditure elasticities of chicken seem unreasonably high. This motivates the imposition of restrictions on the expenditure elasticities discussed next.

Table 1. Expenditure elasticities required to minimize the taste changes in U.S. meat demand

Year	Beef	Pork	Chicken
1972	0.00	0.00	9.62
1973	1.78	0.02	0.00
1974	1.76	0.00	0.00
1975	1.17	0.00	2.85
1976	0.12	1.54	4.52
1977	0.00	0.00	8.40
1978	0.00	0.00	8.62
1979	0.00	0.00	8.85
1980	1.06	0.00	3.33
1981	0.00	3.22	0.00
1982	0.00	0.21	8.06
1983	0.00	0.00	8.47
1984	0.00	0.00	7.35

If we knew the true expenditure elasticities, our test results would be more accurate than the results obtained from using (12). If we attempt to measure these elasticities, however, the model misspecification problem will be reintroduced. To minimize this disadvantage, we now introduce statistical confidence intervals of estimated expenditure elasticities. The hope in doing so is that errors from model misspecification can be minimized. The linear programming problem (12) can be rewritten when the upper and lower bounds of expenditure elasticities are considered:

$$\underset{PTC,\ \phi}{Min} \quad b'\,PTC;$$

$$
\begin{aligned}
s.t.\quad (i)\ & \sum_{i=1}^{n} p_{it}q_{it} - \sum_{i=1}^{n} p_{it}q_{is} \\
& \leq \sum_{i=1}^{n}\sum_{j=2}^{t} a_j \frac{p_{it}q_{ij}}{x_j}\,\epsilon_{ix}^{j} - \sum_{i=1}^{n}\sum_{j=2}^{s} a_j \frac{p_{it}q_{ij}}{x_j}\,\epsilon_{ix}^{j} \\
& + \sum_{i=1}^{n}\sum_{j=2}^{t} p_{it}ptc_{ij} - \sum_{i=1}^{n}\sum_{j=2}^{s} p_{it}ptc_{ij} \\
& \textit{for all } t \textit{ and } s\ ;
\end{aligned}
\tag{13}
$$

$$(ii)\ \sum_{i} w_i^t \epsilon_{ix}^t = 1 \qquad \textit{for all } t\ ;$$

$$(iii)\ \epsilon_{ix}^t \geq 0 \qquad \textit{for all } i \textit{ and } t\ ; \textit{ and}$$

$$(iv)\ \mu^L \leq \phi \leq \mu^U$$

where μ^L and μ^U are vectors of lower bounds and upper bounds, respectively, of estimated expenditure elasticities. The expenditure elasticity of good i, which is derived from (5) or (6), has lower and upper bounds ϵ_{ix}^L and ϵ_{ix}^U,

$$\epsilon_{ix}^L \leq \epsilon_{ix} = \frac{(q_i - q_i^*)x}{aq_i} \leq \epsilon_{ix}^U\,,$$

and thus the compensated demand of good i has a narrower range than does (12):

$$q_i - \frac{aq_i}{x}\epsilon^U_{ix} \le q_i^* \le q_i - \frac{aq_i}{x}\epsilon^L_{ix} \quad \textit{for all } i \ .$$

The expenditure elasticities of meat demand are estimated from the AIDS model as

$$w_{it} = \alpha + \sum_{j=1}^{n} \beta_{ij}\log(p_{jt}) + \beta_i \log(\frac{X}{P^*}) + e_{it} \ ,$$

where P^* is a price index approximated by the Stone geometric index, i.e., $\log(P^*) = \sum_{j=1}^{n} w_i \log(p_i)$, and e_{it} is an error term. The time period estimating the expenditure elasticities is 1960 to 1985.

Table 2 shows the upper and lower bounds of expenditure elasticities at the mean when confidence intervals of 95 percent are used.[8] The U.S. expenditure elasticities for beef range from 1.042 to 1.369, and the chicken elasticities range from 0.392 to 1.374. These elasticities seem more reasonable than those presented in Table 1.

Table 2. Upper and lower bounds of expenditure elasticities at the means

	United States		Canada		Japan		South Korea	
	Lower	Upper	Lower	Upper	Lower	Upper	Lower	Upper
Beef	1.042	1.369	1.149	1.452	0.595	1.780	0.351	0.783
Dairy					1.206	1.789		
Pork	0.462	0.906	0.442	0.860	0.316	0.692	1.394	1.869
Chicken	0.392	1.374	0.529	1.022	1.258	1.543	−0.382	0.760

The results obtained from U.S. meat data using problem (13) are shown in Figure 5, where only the taste changes of beef are represented for graphical convenience.[9] These results indicate a movement of slightly more than 5 pounds against beef. The unreported results for chicken show a positive movement of 2.6 pounds between 1979 and 1984. These numbers are not dramatically different from the results of the first test despite the very restrictive impact of this procedure on the magnitude of the chicken expenditures.

[8] We actually used different lower and upper bounds of expenditure elasticities every year because $\epsilon^i_{ix} = 1 + \beta_i/w_{it}$ and w_{it}'s are different every year.

[9] The results from (13) and (14) indicate a slight taste change in favor of chicken.

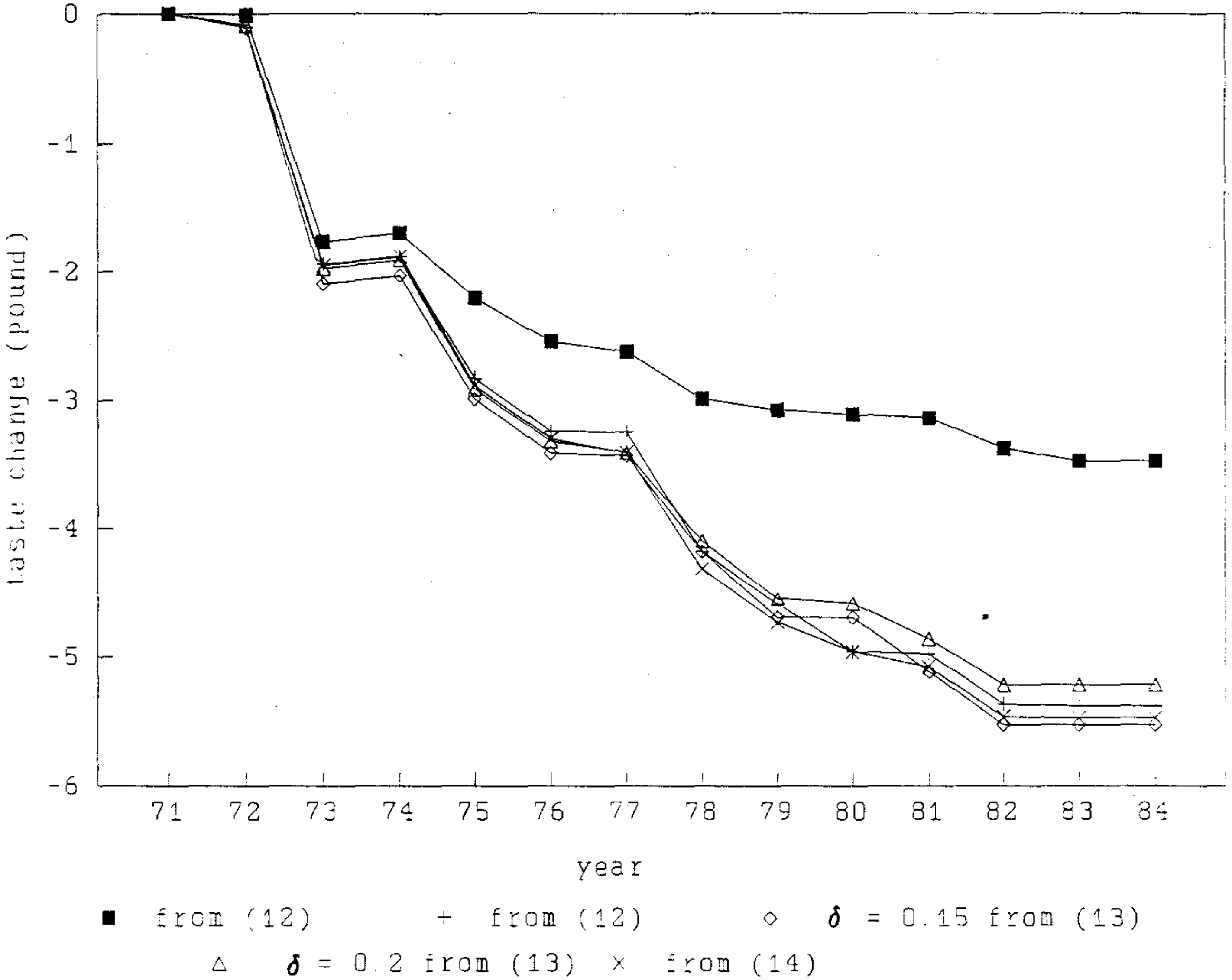

Figure 5. Taste change in U.S. beef demand

A second way of imposing realism on the elasticities from Table 1 is to impose reasonable bounds on how elasticities can change from year to year. For example, Table 1 indicates that the expenditure elasticity for chicken was 9.62 in 1972 and 0.00 in 1973. This result motivates a restriction on the magnitude of the year-to-year changes in expenditure elasticities. This procedure does not depend on any parametric estimates. Suppose that we impose the restriction that the difference of the expenditure elasticities between time t and the previous time $(t - 1)$ for all t less than $\pm\delta$. Then, $|\epsilon_{ix}^{t} - \epsilon_{ix}^{t-1}| \leq \delta$ is used in place of the fourth constraint in (13):

$$\underset{PTC,\phi}{Min} \quad b' \, PTC \tag{14}$$

$$\begin{aligned}
s.t. \quad (i) \; & \sum_{i=1}^{n} p_{it} q_{it} - \sum_{i=1}^{n} p_{it} q_{is} \\
& \leq \sum_{i=1}^{n} \sum_{j=2}^{t} a_j \frac{p_{it} q_{ij}}{x_j} \epsilon_{ix}^{j} - \sum_{i=1}^{n} \sum_{j=2}^{s} a_j \frac{p_{it} q_{ij}}{x_j} \epsilon_{ix}^{j} \\
& + \sum_{i=1}^{n} \sum_{j=2}^{t} p_{it} \, ptc_{ij} - \sum_{i=1}^{n} \sum_{i=2}^{s} p_{it} \, ptc_{ij} \\
& \textit{for all } t \textit{ and } s \; ; \\
(ii) \; & \sum_{i} w_i^t \epsilon_{ix}^{t} = 1 \qquad \textit{for all } t \; ; \\
(iii) \; & \epsilon_{ix}^{t} \geq 0 \qquad \textit{for all } i \textit{ and } t \; ; \textit{ and} \\
(iv) \; & | \, \epsilon_{ix}^{t} - \epsilon_{ix}^{t-1} \, | \leq \delta \qquad \textit{for all } t \textit{ and } i \; .
\end{aligned}$$

The smaller δ is, the larger the magnitude of taste changes. In our applications, $\delta = 0.15$ and 0.2 for all i and t, respectively. That is, the changes of expenditure elasticities at time t are allowed to change from $\epsilon_{ix}^{t-1} - 0.15$ to $\epsilon_{ix}^{t-1} + 0.15$ and from $\epsilon_{ix}^{t-1} - 0.2$ to $\epsilon_{ix}^{t-1} + 0.2$, respectively.

The results obtained for beef from the U.S. data are also shown in Figure 5. The lines "$+-+-+$" and "$\diamond$-$\diamond$-$\diamond$" show the taste changes obtained by using $\delta = 0.15$ and $\delta = 0.2$ in problem (14). The taste changes in both cases are almost identical.

One could also place the year-to-year restriction on the estimated elasticities of the second procedure; however, this procedure does not change the results in any significant way.

For the United States, one may conclude that some consistent bias has existed against beef. The cumulative effect of this bias has been somewhere between 3.5 and 5 pounds over the study period. We cannot tell if this inconsistency is attributable to some systematic error in the data, i.e., a gradual underreporting of the amount of fat cut off beef, or because consumer preferences have in fact moved against beef. The magnitude of this bias seems small, however, when compared with the more than 15-pound decrease in consumption observed between 1976 and 1984.

The U.S. results demonstrate the ability of the new method to detect relatively small changes in preferences. Given the standard errors usual in parametric work, it is unlikely that one could ever provide convincing evidence of a one- or two-pound per capita change in preferences. Also, neither Chalfant and Alston nor Cox and Chavas detected any taste change when nonparametric methods were used. This ability to detect slight changes may allow for an accurate measurement of the positive effects (if any) of the enormous generic campaigns for pork and poultry that occurred in the 1980s.

Figures 6 through 8 show results from (12) and (13) for Canada, Japan, and South Korea, respectively. These results are expressed in kilograms per capita. The results for Canada are very similar to those for the United States, with a maximum shift against beef of 3 kilograms and a move in favor of poultry of almost 3 kilograms. The Canadian results indicate a slight movement away from pork that occurs until almost 10 years later in the United States. One might hypothesize that because Canadian meat data are collected by a different agency and in a different manner than in the United States while consumers in both countries receive similar nutritional information, the source of the inconsistency in both countries is consumer- rather than data-driven.

The results for Japan show a positive movement toward native Japanese, or Wagyu, beef and a negative movement against Japanese dairy beef and imported beef. The magnitude of these changes is very small but represents a significant proportion of consumption. (In 1984, Japanese consumers ate 1.089 kilograms of Wagyu beef and 2.751 kilograms of dairy beef, respectively.) There is evidence of a slight shift away from pork in Japan while all the chicken data were consistent with preferences. Wahl, Hayes, and Williams (1991) report that Japanese farmers replaced Wagyu animals with tractors

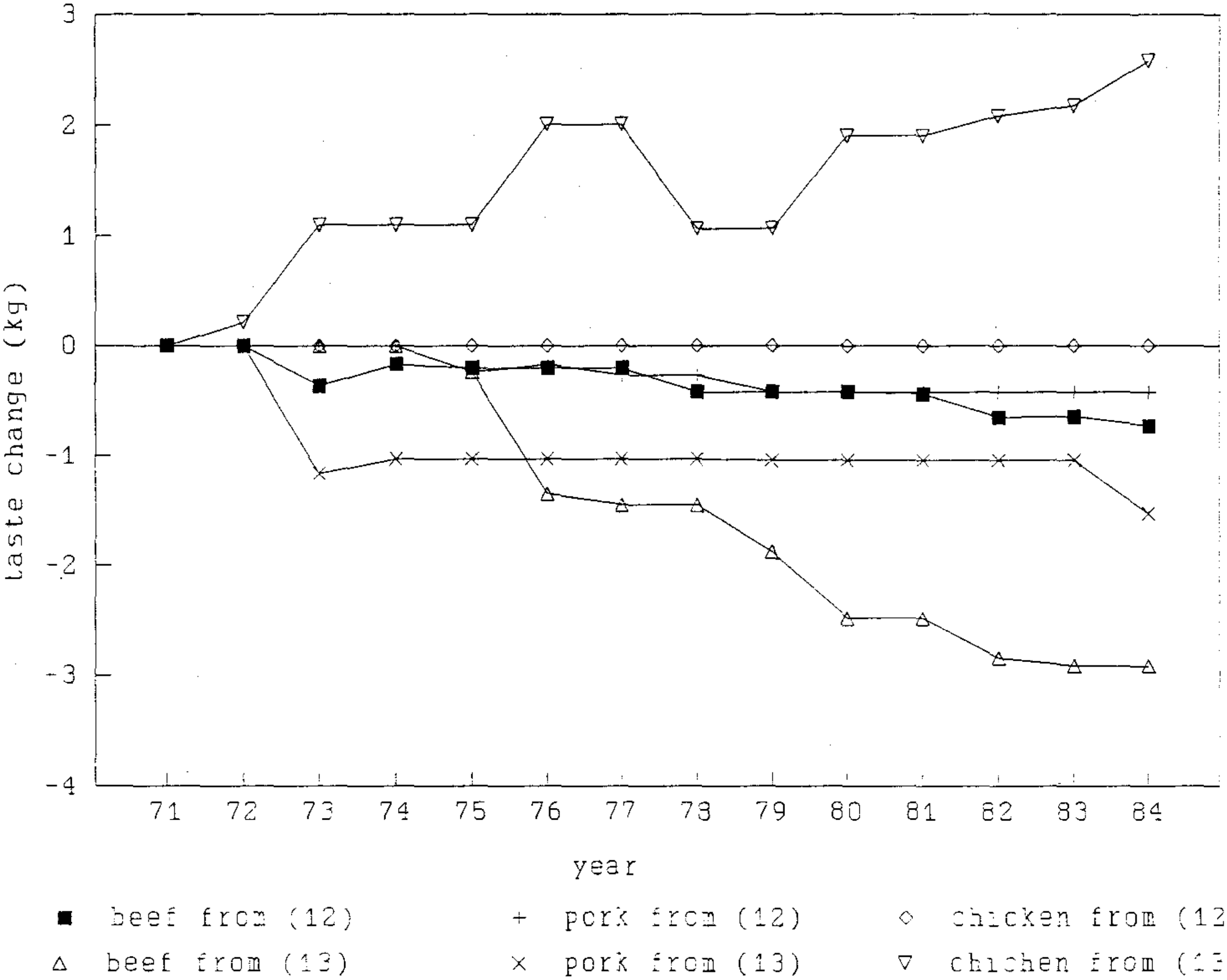

Figure 6. Taste change in Canadian meat demand

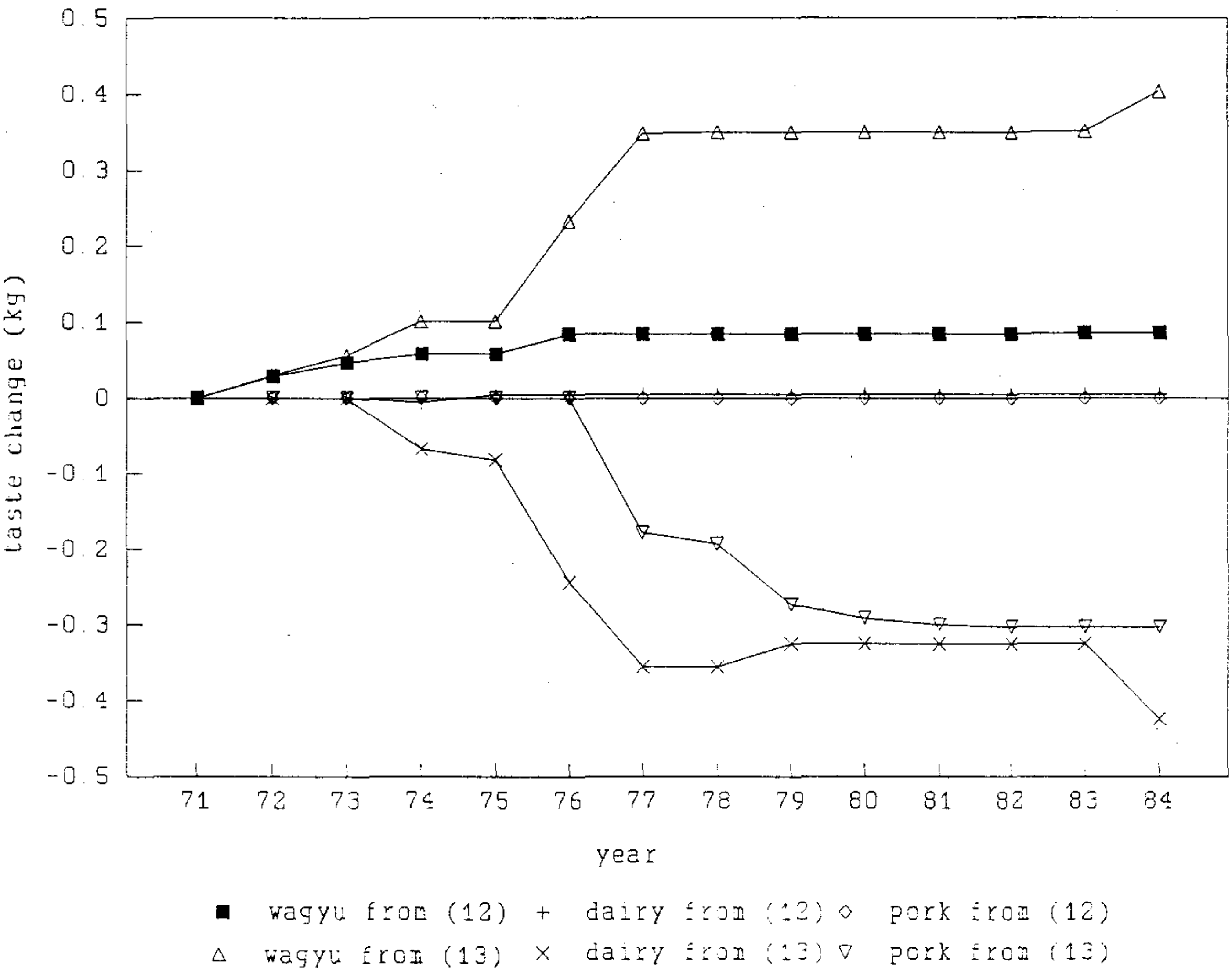

Figure 7. Taste change in Japanese meat demand

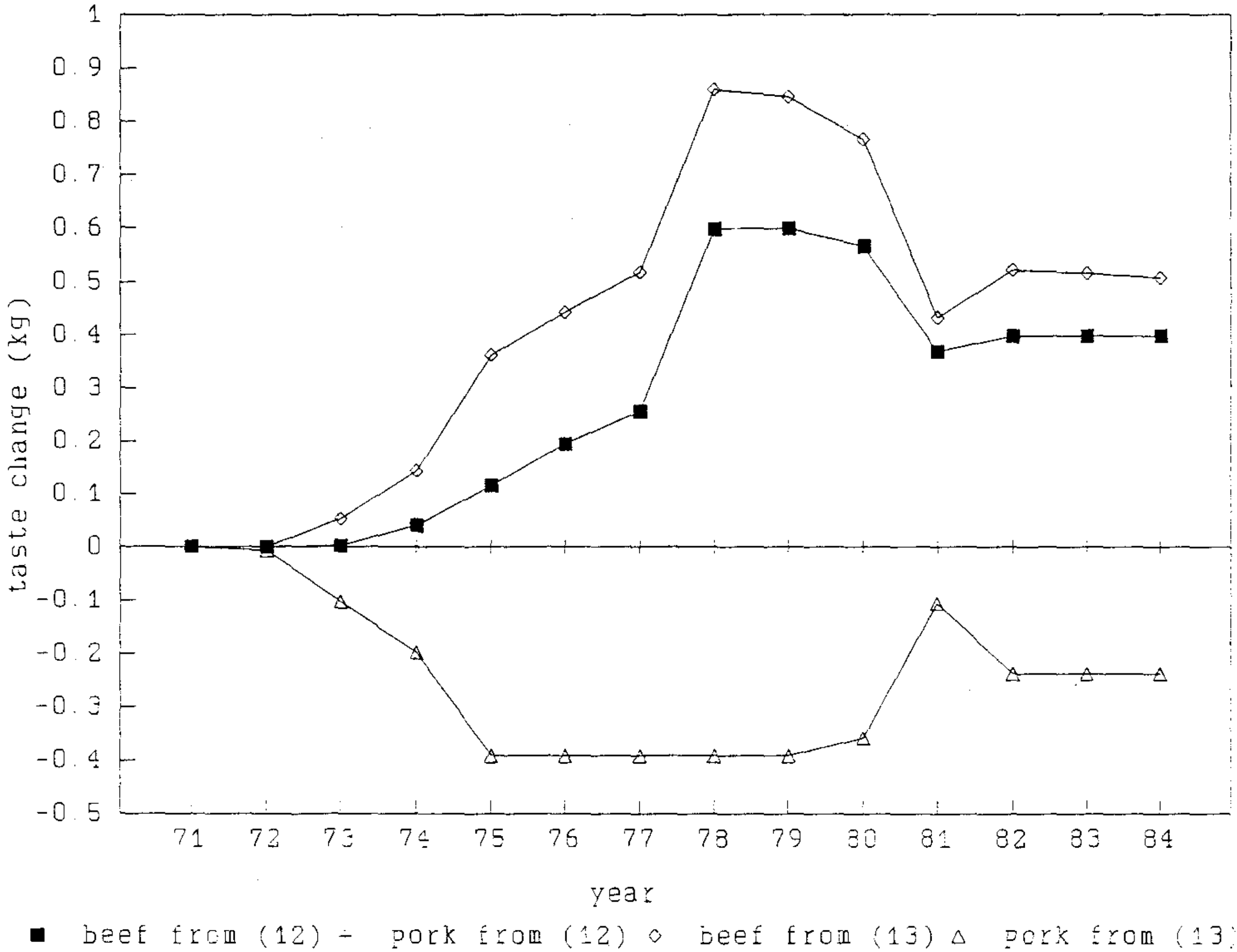

Figure 8. Taste change in Korean meat demand

in the early 1970s. This change means that the quality of Wagyu beef would have improved considerably during this period. In the Japanese government statistics, data for Wagyu animals do notdifferentiate between retired draft animals and younger custom-fed beef animals. It seems likely, therefore, that the source of the inconsistency in Japan was data-driven rather than consumer-driven.

The South Korean results indicate a positive movement toward beef up to 1978, followed by a slight decrease to 1984. A slight movement against pork occurred between 1972 and 1976, but this was almost reversed in 1981. Again, no violations in the chicken data were detected.

Conclusions

This paper presents a way of detecting violation of consumer preference theory that has some of the advantages of existing parametric and non-parametric methods. The proposed method does not require any subjective input on behalf of the modeler and is therefore less subject to pretesting and data mining. The new method has the ability to detect slight violations in preferences, even when the budget constraint has shifted out, a feature that has not been found in the non-parametric models that have been presented to date.

The model was used to examine meat demand data for the United States, Canada, Japan, and South Korea. The results indicate a shift away from beef in the United States and Canada, while the opposite may have been the case in Japan and South Korea. Smaller negative shifts have occurred against pork in all four countries. U.S. and especially Canadian consumers appear to have moved toward chicken, whereas Japanese and South Korean consumers have remained neutral.

REFERENCES

Agricultural Cooperative Federation. *Agricultural Cooperative Yearbook.* Seoul, South Korea, various issues.

Alston, J.M., and J.A. Chalfant. "Can We Take the Con Out of Meat Demand Studies?" *Western Journal of Agricultural Economics*, forthcoming.

____________________. "Unstable Models from Incorrect Forms." *American Journal of Agricultural Economics*, forthcoming.

____________________. "Effects of Functional Form Choices on Tests for Structural Change in Demand." University of California, Berkeley, 1991c.

Chalfant, J. "A Global by Flexible, Almost Ideal Demand System." Journal of Business and Economic Statistics 5(1987):233-42.

Chalfant, J., and J. Alston. "Accounting for Changes in Taste." *Journal of Political Economy* 96(1988):391-410.

Chavas, J.P., and T.L. Cox. "A Nonparametric Analysis of Productivity: The Case of U.S. and Japanese Manufacturing." *American Economic Review* 80(1990):450-64.

Cox, T.L., and J.P. Chavas. "A Nonparametric Analysis of the Structure and Stability of Preferences." Department of Agricultural Economics, University of Wisconsin-Madison, April 1990.

____________________. "Nonparametric Demand Analysis: A Duality Approach." Department of Agricultural Economics, University of Wisconsin-Madison, Staff Paper No. 265, January 1987.

Leamer, E.E. "Let's Take the Con Out of Econometrics." *American Economic Review* 73(1983):31-43.

National Livestock Cooperatives Federation. *Materials on Price, Demand, and Supply for Livestock Products*. Seoul, South Korea, various issues.

Van Kooten, G.C. "The Economic Impacts on Consumers of Government Intervention in the Poultry and Egg Sectors: A Comparison of Alternative Welfare Measures." Agriculture Canada, Working Paper 5/87, March 1987.

Varian, H.R. *Microeconimc Analysis*. New York: Norton, 1984.

Wahl, T.I., and D.J. Hayes. "Demand System Estimation with Upward-Sloping Supply." *Canadian Journal of Agricultural Economics* 38(March 1990):107-22.

Wahl, T.I., D.J. Hayes, and G.W. Williams. "Dynamic Adjustment in the Japanese Beef Industry under Beef Import Liberalization." *American Journal of Agricultural Economics* 73(1991):118-32.