Import Demand for Disaggregated Fresh Fruits in Japan

Troy G. Schmitz and James L. Seale, Jr.

Using annual Japanese fresh fruit import data from 1971 to 1997, this study analyzes the import patterns of Japan's seven most popular fresh fruits by implementing and testing a general differential demand system that nests four alternative import demand specifications. When tested against the general system using the five-good case (bananas, grapefruits, oranges, and lemons and aggregating pineapples, berries, and grapes), the analysis rejects the Almost Ideal Demand System and National Bureau of Research specifications but does not reject Rotterdam and Central Bureau of Statistics models. When estimated using the six-good case (bananas, grapefruits, oranges, lemons, and pineapples and aggregating berries and grapes), the analysis rejects all specifications except the Rotterdam model.

Key Words: Almost Ideal Demand System, consumer demand, fruit, import demand, Japan, Rotterdam

JEL Classifications: C3, F1, Q0

Extremely high production costs for most agricultural products and the liberalization of several formal barriers to trade as a result of the World Trade Organization (WTO) put Japanese producers under considerable competitive pressures. As the number of Japanese producers has steadily declined (a 14% reduction in 1998 compared with 1990), Japan has become increasingly dependent upon agricultural

imports. As a result, Japan is the world's largest importer of agricultural products, importing $33 billion worth of agricultural products in 1999 (USDA/ERS).

Japan has made several steps toward deregulation in the fruit industry since 1988. These include (1) import quota reductions for fresh orange imports from the United States in 1988, (2) tariff reductions for grapefruit and lemons in 1989, (3) removal of import quotas for fresh oranges in 1991, (4) lifting the import ban on apple imports from New Zealand in 1993, and (5) lifting the import ban on apples from the United States in 1994. Furthermore, in July 1999, Japan adopted a new philosophy on agricultural policy, choosing to focus on national food security, multifunctionality, and less trade-distorting policies. However, the objective of food security is still carried out by not allowing the share of imports of many agricultural products to exceed 60% of domestic caloric intake. Furthermore, bound and applied tariffs are still significant for many fruit

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Table 1. Japanese Fresh Fruit Imports, 1997

Fresh Fruit	Value Billion Yen	Quantity Million Tons
Bananas	52.9	885.5
Grapefruit	31.0	283.8
Oranges	18.1	171.3
Lemons	17.7	89.4
Pineapples	5.5	96.1
Berries	4.4	4.8
Grapes	2.2	7.3

imports, regardless as to whether the exporting country is a member of the WTO or not.

The Japanese fresh fruit market is an important component of U.S. agricultural exports. Fluctuations in prices caused by variable market conditions and changes in Japanese import policies have caused U.S. exports of fresh fruit to Japan to fluctuate. The main objective of this article is to empirically estimate the sensitivity of Japanese fresh fruit imports to changes in Japanese income levels and import prices. The Japanese fresh fruit market was chosen because it is a relatively important export market for producers in the southern United States. A list of the major types of fresh fruit imported by Japan in 1997, in terms of both value and quantity, is provided in Table 1. Notice that bananas, grapefruit, oranges, and lemons are the most important fresh fruit imports from the perspective of Japanese consumers. Although grapefruit, oranges, and lemons are important in terms of U.S. agricultural exports, bananas are not. However, it may be that bananas act as a substitute for grapefruit, oranges, and lemons from the perspective of Japanese consumers. Hence, all of these fruits should be considered when attempting to estimate the response of Japanese consumers to changes in relative import prices.

Although the fresh fruit market has become increasingly important in terms of its contribution to the total value of U.S. agricultural exports, there are relatively few empirical demand studies that have focused on the major U.S. markets for disaggregate fresh fruit commodities. Most import-demand studies of related products found in the literature have fo-

cused on the demand for aggregate groupings of fruits or vegetables. For example, Sarris (1981, 1983) estimates income and price elasticities of demand for five broad categories of fruits and vegetables (fresh fruits, dried fruits, processed fruits, fresh vegetables, and processed vegetables) in the European Union. Sparks estimated a world trade model for vegetables in which all vegetables and related products are combined into one category.[1] Hunt estimated import demand for 36 disaggregate fruit and vegetable products from Mediterranean countries by the European Union under the assumption that demand is a linear function of per capita income and that market shares are constant. Two studies (Atkin and Blanford; Roberts and Cuthbertson) examined the import demand for fresh apples in the United Kingdom, but apples from the United States were not included in the analysis.

Studies that have estimated demand for aggregate groupings of fresh and processed fruits and vegetables have been limited in the sense that income and price responses may differ markedly among disaggregate products (e.g., apples, oranges, or lemons). They have not taken into account the effect that demand for one good has on that of other similar goods either through a general or specific price substitution effect. Studies that have analyzed the domestic or import demand for fresh and processed fruits and vegetables at a disaggregate level in a systemwide approach have appeared in the literature only recently.

Four recently published studies have addressed the issue of aggregate fresh fruit demand. Lee, Seale, and Jierwiriyapant analyzed the relationships among major suppliers of citrus juices in Japan using a Rotterdam import allocation model. They showed that Japanese demand for imports of fresh grapefruit from the United States is affected by banana and pineapple imports and that the Japanese im-

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[1] Other studies of the vegetable trade do not typically use rigorous empirical estimation techniques and are based on more descriptive or institutional approaches (e.g., Davis and Seale; Fairchild et al.; Kobayashi 1989a,b; Mackintosh; Montegaud and Lauret; Seale; Seale, Davis, and Mulkey).

port demand for U.S. citrus juice is affected by Brazilian and Israeli export competition. Seale, Sparks, and Buxton also applied a Rotterdam model to study the import demand for fresh apples in Canada, Hong Kong, Singapore, and the United Kingdom. It was shown that, except for the case of U.K. imports from Australia, an increase in the total expenditure on apple imports in each of the major apple-importing countries would increase apple exports in each of the major exporting markets. It was also shown that a 1% increase in the expenditure on fresh apple imports in Hong Kong, Singapore, and the United Kingdom would increase imports of U.S. fresh apples by more than 1% in each of these countries. Lee, Brown, and Seale used a nested approach to analyze Canadian fresh fruit and juice import demand for the period from 1960 through 1987. The approach chose between the Rotterdam demand specification and an income-variant differential demand specification developed by Keller and Driel and Clements. Results indicated that if the total expenditure on aggregate Canadian imports of fresh fruit and juices increase, expenditure shares of oranges and apples increase. Furthermore, oranges and grapefruits are substitutes for apples. Hence, an increase in the price of fresh apples would increase the total consumption of citrus, thereby increasing Canadian citrus imports.

Theoretical Model

Empirical demand relationships are estimated under five different econometric specifications. These different specifications are developed under a systemwide approach to consumer demand with multistage budgeting. With two exceptions, the empirical analysis relies on the differential demand system developed by Barten (1964) and Theil (1965).

The most popular estimable demand system that results from the differential approach is known as the Rotterdam model. However, the Rotterdam model is only one particular parameterization adapted from Theil and Barten's work. The Central Bureau of Statistics (CBS) model, developed by Keller and van

Driel and also by Clements, is an alternative parameterization of the differential approach based on Working's model. It assumes that the budget share allocated to each commodity group is a linear function of the logarithm of income, whereas the Rotterdam model assumes constant marginal shares. In addition to the differential models, empirical estimates of Japanese fresh fruit demand are obtained for the time-series version of the Almost Ideal Demand System (AIDS) developed by Deaton and Muellbauer and the AIDS income-variant National Bureau of Research (NBR) specification developed by Neves. Furthermore, differential versions of these four demand specifications are nested into a general model (Barten 1993). The results of their empirical application to disaggregate Japanese fresh fruit imports are compared and contrasted.

The Rotterdam Model is derived by starting with utility maximization subject to the budget constraint, which can be written as

$$(1) \quad \max U(q) \quad \text{s.t.} \quad \sum_i (p_i q_i) = M,$$

where $U(q)$ is utility as a function of the consumption of a vector of goods (q), M is total income, p_i is the price of the ith good, and q_i is the quantity of the ith good. However, before utility is maximized, the differential approach to demand system analysis proceeds by totally differentiating the budget constraint, which yields

$$(2) \quad dM = \sum q_i dp_i + \sum p_i dq_i.$$

Dividing Equation (2) through by income (M), multiplying and dividing the first term on the right-hand side (RHS) by p_i, and multiplying and dividing the second term on the RHS by q_i yields

$$(3) \quad \frac{dM}{M} = \sum \left(\frac{p_i q_i}{M}\right)\left(\frac{dp_i}{p_i}\right) + \sum \left(\frac{p_i q_i}{M}\right)\left(\frac{dq_i}{q_i}\right).$$

If you let $w_i = (p_i q_i)/M$ be the budget share of the ith good and make use of the fact that, for any variable X, $dX/X = d(\ln X)$, then Equation (3) can be rewritten as

$$(4) \quad d(\ln M) = \sum w_i \, d(\ln p_i) + \sum w_i \, d(\ln q_i).$$

Using the definitions of both the Divisia price index $[d \ln P = \sum w_i d(\ln p_i)]$ and the Divisia volume index $[d \ln Q = \sum w_i d(\ln q_i)]$, Equation (4) becomes simply

$$(5) \quad d \ln M = d \ln P + d \ln Q.$$

Now, because all terms are in natural logarithms, Equation (5) is (theoretically) exactly equivalent to

$$(6) \quad d[\ln(M/P)] = d \ln Q.$$

Equation (6) depicts the fact that the natural logarithm of the change in income deflated by the price index is equal to the Divisia volume index. Hence, the two can be used interchangeably for theoretical purposes.[2]

Using the above differential relationship for the budget constraint in combination with the solution to Barten's (1964) fundamental matrix,[3] utility maximization eventually leads to the following specification, known as the Rotterdam model (with time subscripts omitted for convenience):

$$(7) \quad w_i \, d \ln q_i = \theta_i d \ln Q + \sum_j \pi_{ij} \, d \ln p_j,$$

$$i = 1, 2, \ldots, n,$$

where $w_i = (w_{it} + w_{i,t-1})/2$ represents the av-

erage value share for commodity i with subscript t standing for time, $d \ln q_i = \ln(q_{it}/q_{i,t-1})$ is the natural logarithm of the change in the consumption level for commodity i, $d \ln p_i = \ln(p_{it}/p_{i,t-1})$ is the natural logarithm of the change in the price for commodity i, and $d \ln Q$ is the Divisia volume index for the change in real income as in Equation (6).[4]

The solution to Barten's fundamental matrix also yields the following relationships for the demand parameters

$$(8) \quad \theta_i = p_i(\partial q_i/\partial M) \quad \text{and} \quad \pi_{ij} = (p_i p_j/M)s_{ij},$$

where $s_{ij} = \partial q_i/\partial p_j + q_j \partial q_i/\partial M$, M is total outlay or the budget, and s_{ij} is the (i, j)th element of the Slutsky substitution matrix. The parameter θ_i is the marginal budget share for commodity i and π_{ij} is a compensated price effect. Because of the strict theoretical constructs that the Rotterdam model adheres to, the following constraints of demand theory must be directly applied to its parameters:

$$(9) \quad \text{Adding-up} \quad \sum_i \theta_i = 1,$$

$$\sum_i \pi_{ij} = 0;$$

$$(10) \quad \text{Homogeneity} \quad \sum_j \pi_{ij} = 0; \quad \text{and}$$

$$(11) \quad \text{Slutsky symmetry} \quad \pi_{ij} = \pi_{ji}.$$

The Rotterdam model is a particular parameterization of a system of differential demand equations where the demand parameters, θ_i and π_{ij}, are assumed to be constant. However, there is no strong prior reason that the θ_i and π_{ij} should be held constant. An alternative parameterization is based on Working's Engel model,

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[2] In empirical application, it is important to replace $d(\ln M/P)$ with $d(\ln Q)$ so as to ensure that the adding-up conditions are met and that the sum of the error terms over all i equations equals zero. Theil (1971, p. 332) proved that the empirically based $d(\ln M/P)$ and $d(\ln Q)$ differ only by a term of third-order smallness.

[3] The development of the Lagrangian technique for solving this utility maximization problem eventually leads to what has become known as Barten's fundamental matrix (Barten 1964, pp. 2–3). Essentially, this matrix makes direct use of the Hessian to formulate a set of equations that are then solved to yield the Rotterdam specification. Although a thorough discussion is beyond the scope of this article (see Theil 1980 for a more elegant explanation), it is interesting to note that the reason that one always imposes symmetry on the Rotterdam system is because Young's theorem of derivatives under continuous functions dictates that the Hessian be symmetric.

[4] When performing empirical analysis using the Rotterdam specification, $d \ln X_i$ must be computed as the difference between the logarithm of the value of X in the current year and the logarithm of the value of X in the previous year for any variable X. Hence, because the differential approach uses results from the total differentiation of the budget constraint, theory dictates the use of log differences in applications of the Rotterdam model. When one follows this approach using data over time, the first observation necessarily gets dropped.

(12) $\quad w_i = \alpha_i + \beta_i \ln M, \qquad i = 1, 2, \ldots, n.$

As the sum of the budget shares is unity, it follows from Equation (12) that $\Sigma \alpha_i = 1$ and $\Sigma \beta_i = 0$. To derive the marginal shares implied by Working's model, one multiplies Equation (12) by M and then differentiates with respect to M, which results in

(13) $\quad \partial(p_i q_i)/\partial M = \alpha_i + \beta_i(1 + \ln M) = w_i + \beta_i.$

Hence, under Working's model, the ith marginal share differs from the corresponding budget share by β_i; because the budget share is not constant with respect to income, neither is the associated marginal share. The expenditure elasticity corresponding to Equation (13) is

(14) $\quad \eta_i = 1 + \beta_i/w_i.$

This expression indicates that a good with positive (negative) β_i is a luxury (necessity). Because the budget share of a luxury increases with income (prices remaining constant), it follows from Equation (14) that increasing income causes the η_i for such a good to fall toward one. The income elasticity of a necessity also declines with increasing income under Equation (14). Accordingly, as the consumer becomes more affluent, luxury and necessity goods become less luxurious under Working's model, a plausible outcome. If $\beta_i = 0$, however, the good is unitary elastic and the budget share will not change in response to income changes (again, with prices held constant).

Replacing θ_i in Equation (7) with Equation (13) and rearranging terms, one obtains

(15) $\quad w_i(d \ln q_i - d \ln Q)$

$\qquad = \beta_i \, d \ln Q + \sum_j \pi_{ij} \, d \ln p_i.$

where β_i and π_{ij} are constant coefficients (Clements; Keller and van Driel). Equation (15) will be referred to as the CBS model following Keller and van Driel.

The AIDS model, another specification, is specified as

(16) $\quad w_i = \alpha_i + \sum_j \gamma_{ij} \ln p_j + \beta_i \ln(M/P),$

where P is a price index defined by $\log P = \alpha_0 + \Sigma \alpha_k \ln p_k + \frac{1}{2} \Sigma_k \Sigma_l \gamma_{kl} \ln p_k \ln p_l.$

The adding up restriction requires that

$$\sum_i \alpha_i = 1, \qquad \sum_i \beta_i = 0, \qquad \sum_i \gamma_{ij} = 0.$$

Homogeneity is satisfied if and only if $\Sigma_j \gamma_{ji} = 0$, and symmetry is satisfied provided that $\gamma_{ij} = \gamma_{ji}$.

By approximating P by Stone's price index and the logarithmic change in Stone's price index by the Divisia price index, $\Sigma_i w_i \, d \ln p_i$, Equation (16) can be expressed in differential form (Barten 1993; Deaton and Muellbauer),

(17) $\quad dw_i = \beta_i \, d \ln Q + \sum_j \gamma_{ij} \, d \ln p_j.$

As shown by Barten (1993), $\beta_i = \theta_i - w_i$, and $\gamma_{ij} = \pi_{ij} + w_i \delta_{ij} - w_i w_j$, where δ_{ij} is the Kronecker delta equal to unity if $i = j$ and zero otherwise. Note that the CBS system has the AIDS income coefficients β_i and the Rotterdam price coefficients π_{ij}. Also, if all units of analysis face the same prices, both the CBS and AIDS collapse to the simple Working's model.

Another alternative model, the NBR model (Neves), can be derived by substituting $\theta_i - w_i$ for β_i in Equation (17) so that it has the Rotterdam income coefficients but the AIDS price coefficients (Barten 1993). Specifically, the NBR model is

(18) $\quad dw_i + w_i \, d \ln Q = \theta_i \, d \ln Q + \sum_j \gamma_{ij} \, d \ln p_j,$

and the NBR and the CBS models can be considered as income-response variants of the Rotterdam and AIDS models, respectively.

These four models are not nested, but, following Barten (1993), a general model can be developed which nests all four models. Specifically, the general model is

(19) $w_i\, d \ln q_i$

$$= (d_i + \delta_1 w_i)\, d \ln Q + \sum_j e_{ij}\, d \ln p_j$$

$$+ \delta_1 w_i\, d \ln Q - \delta_2 w_i (d \ln p_i - d \ln P);$$

$$i = 1, 2, \ldots, n,$$

where δ_1 and δ_2 are two additional parameters to be estimated. Note that Equation (19) becomes the Rotterdam model when both δ_1 and δ_2 are restricted to be zero, the CBS model when $\delta_1 = 1$ and $\delta_2 = 0$, the AIDS model when $\delta_1 = 0$ and $\delta_2 = 1$, and the NBR model when $\delta_1 = 1$ and $\delta_2 = 1$. The demand restrictions on Equation (19) are

Adding-up $\sum_i d_i = 1 - \delta_1$ and

$$\sum_i e_{ij} = 0;$$

Homogeneity $\sum_j e_{ij} = 0;$ and

Symmetry $e_{ij} = e_{ji}.$

Data Description

Import expenditure data regarding the volume and the value (in Japanese Yen) of all major types of fresh fruit imported by Japan were collected from the United Nations Trade Data Tape. The United Nations Trade Data Tape contains annual data from 1971 through 1997 and aggregates imports from all source countries for each individual good. Because of the massive effort in reporting, collecting, confirming, and finalizing these data sets for all countries involved, the data contained in these data sets lag anywhere from 2 to 4 years. Hence, because of data limitations, our period of analysis ends in 1997. The import expenditure shares in Table 2 provide a summary of the United Nations data regarding Japanese fresh fruit imports. The seven major types of fresh fruit imported by Japan in a typical year are, in order of value, (1) bananas, (2) grapefruit, (3) oranges, (4) lemons, (5) pineapples, (6) berries, and (7) grapes. The theoretical models presented above and estimated below attempt to explain the changes in relative import expenditure shares over time as a function

of the change in relative prices of the different types of fruit over time.[5]

Estimation Procedure

Because of the adding-up restrictions, the full $n \times n$ matrices of all five systems are singular (n is the number of goods). Barten (1969) proved that, by omitting one equation and estimating the $n - 1$ system of equations, the parameter estimates are invariant to which equation is omitted. Hence, we drop the other fruit equation and estimate all five systems with iterative seemingly unrelated regressions (SUR) that iterates to the maximum likelihood (ML) estimator. This is accomplished by using the LSQ command in Time Series Processing (TSP) version 4.3.

We also tested all five systems for autocorrelation of degree one (AR1) in the error terms by transforming the data with the Prais-Winston transformation and constraining the AR1 parameter, rho, to be the same in all n. Because the Jacobian term is no longer equal to one (or the log of the Jacobian term is not equal to zero), iterative SUR is not ML (Theil, Chung, and Seale). To obtain ML estimates of rho and all the other parameters from the AR1-specified models, we used the Hildreth-Lu ML procedure. The log-likelihood ratio test was used to test whether the rho parameter was statistically equal to zero; the unrestricted model is AR1, whereas the restricted model does not have autocorrelation. In all cases, AR1 is soundly rejected. For example, for the Rotterdam system with homogeneity and symmetry imposed, the ML estimate of rho is .01, and the chi-square statistic is only .01, whereas the critical value at the 95% confidence level is 3.84. Because the Rotterdam system as

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[5] Note that apple imports by Japan can at times be significant. However, because of Japanese import policy and the prevalence of Fuji brand apples grown in relatively large quantities in Japan, New Zealand was not allowed to export apples to Japan until 1993, and the U.S. ban on apple exports to Japan was lifted in 1994. Furthermore, when the historic data were originally purchased (at a significant cost) from the United Nations, funding for the project did not include the cost of purchasing apple data. Hence, apples are not included in this study.

Table 2. Import Expenditure Shares of Seven Types of Japanese Fresh Fruit, 1970–1997

Year	Bananas	Grapefruit	Oranges	Lemons	Pineapples	Berries	Grapes
1970	.82	.00[a]	.01	.14	.03	.00	.00
1971	.76	.02	.01	.16	.04	.00	.00
1972	.63	.14	.02	.16	.04	.00	.00
1973	.55	.16	.03	.21	.04	.00	.00
1974	.51	.20	.03	.22	.03	.00	.00
1975	.55	.20	.04	.17	.04	.00	.00
1976	.53	.20	.04	.17	.05	.00	.01
1977	.50	.22	.03	.17	.07	.00	.00
1978	.44	.17	.09	.21	.08	.00	.01
1979	.41	.19	.09	.22	.08	.01	.00
1980	.43	.18	.10	.19	.09	.01	.01
1981	.42	.21	.12	.16	.09	.01	.00
1982	.44	.18	.14	.15	.08	.01	.01
1983	.44	.20	.12	.16	.07	.01	.01
1984	.45	.16	.14	.16	.06	.02	.01
1985	.48	.12	.14	.16	.07	.02	.01
1986	.48	.16	.12	.13	.08	.01	.01
1987	.43	.18	.14	.14	.08	.02	.01
1988	.44	.19	.13	.13	.07	.02	.02
1989	.42	.22	.13	.13	.07	.02	.01
1990	.44	.17	.15	.13	.06	.02	.02
1991	.42	.23	.12	.14	.05	.02	.02
1992	.46	.22	.14	.10	.05	.02	.02
1993	.45	.20	.15	.12	.05	.02	.02
1994	.39	.24	.17	.11	.05	.03	.02
1995	.38	.24	.16	.12	.04	.03	.02
1996	.40	.25	.15	.12	.04	.03	.02
1997	.40	.24	.14	.13	.04	.03	.02
Mean	.48	.18	.10	.15	.06	.01	.01

[a] The value .00 does not indicate that the number is exactly equal to zero but that the number is rounded to .00; the number is positive but less than .005.

well as the other four systems fit the data in log difference, this is not surprising.[6]

Empirical Results for Five Goods

Testing Restriction, Choice of Functional Forms, and Goodness-Of-Fit

In this section, we present empirical estimates of behavioral relationships that partially explain Japanese import patterns for different

[6] At the recommendation of a reviewer, we also tested for the stationarity of the log-differenced data with Dickey-Fuller tests. For all variables in log differences, the Dickey-Fuller tests strongly rejected unit roots, indicating that the data in log differences are all stationary.

fruits. First, we estimate five unrestricted demand systems including the unrestricted general demand system, Equation (14), and then constrain the five systems by imposing homogeneity and then symmetry. The log-likelihood values associated with each of these demand systems are provided in Table 3. The numbers in parentheses are equal to the number of free parameters. The log-likelihood ratio test (LRT) statistic is LRT $= -2[\log L(\theta^*) - \log L(\theta)]$, where θ^* is the vector of parameter estimates with the restrictions imposed, θ is the vector of parameter estimates without the restrictions, and $\log L(\cdot)$ is the log value of the likelihood function. This statistic must be compared with a critical value from a $\chi^2(q)$ distribution, where q is the number of restric-

Table 3. Log-likelihood Values and Alternatives for Five Japanese Fresh Fruit Imports Demand Systems, 1971–1997

Restriction	Model				
	General	Rotterdam	CBS	AIDS	NBR
Unrestricted	311.6 (26)[a]	311.3 (24)	309.7 (24)	305.9 (24)	306.0 (24)
Homogeneity	310.2 (22)	310.0 (20)	307.8 (20)	303.7 (20)	304.4 (20)
Symmetry[b]	305.2 (16)	304.75 (14)	302.9 (14)	299.5 (14)	299.5 (14)

[a] Number of free parameters for each model is in parentheses.

[b] Symmetry is the case in which both homogeneity and symmetry are jointly imposed.

tions imposed (Harvey, pp. 160–166). For example, the unrestricted log-likelihood value for the general model is 311.6. The restricted log-likelihood value for the test of homogeneity in the general model is 310.2. Hence, LRT = $-2(310.2 - 311.6) = 2.8$. The critical value for the test has degrees of freedom equal to the difference in the number of free parameters between the general unrestricted model in Equation (26) and the number of free parameters in the general model with homogeneity imposed in Equation (22). The critical value for this case is a χ^2 value with 4 degrees of freedom. At a 95% level of significance, this critical value is 9.348. Hence, because the LRT statistic is not in the rejection region, we fail to reject the hypothesis of homogeneity. If one performs this comparison for all different combinations of likelihood ratio values in Table 3 (implying different critical values for each comparison, since the degrees of freedom differ), the results indicate that we fail to reject any of the two economic constraints, homogeneity or symmetry, with any of the five models at a 95% level of significance.

Log-likelihood tests were also undertaken between the general model, with homogeneity and symmetry imposed, and each of the other four models (same restrictions) that are nested within the general demand system. When performing cross-model comparisons, the critical χ^2 value always has 2 degrees of freedom and is equal to 5.991 at the 95% level of significance. The Rotterdam model is not rejected at the 95% confidence level, whereas the CBS model is not rejected at the 90% level of significance. The AIDS and NBR models are both strongly rejected at the 90% level of significance.

Further evidence on the fit of the systems is provided by calculating a system-wide R^2 (McElroy).[7] The measure is

$$R_s^2 = 1 - \frac{1}{1 + W^*/(T - K)(n - 1)},$$

where T is the number of observations, K is the number of estimated parameters in each equation, n is the number of equations in the full system, and W^* is a small-sample corrected Wald test statistic under the hypothesis that all estimated parameters in the system are zero. It is interesting to note that the R_s^2s for the general, Rotterdam, CBS, AIDS, and NBR systems are .52, .99, .69, .49, and .98, respectively. This result seems to suggests that, for these data, the constant marginal shares of the Rotterdam and NBR systems have higher explanatory power than those based on Working's model.

Parameter Estimates

Individual parameter estimates for the General, Rotterdam, and CBS models as estimated using the procedure discussed in the previous section are provided in Table 4. We did not include parameter estimates for the AIDS or NBR model because these models were rejected at a 90% level of significance. In the general model, the expenditure coefficient for grapefruit is significant at $\alpha = .05$, whereas the expenditure coefficients for oranges and others are statistically significant at $\alpha = .10$. Neither d_1 nor d_2 are significantly different

[7] Single-equation measures of R^2 are not appropriate measures of the goodness-of-fit of a system of equations (Bewley; Buse; Glahn).

Table 4. Japanese Fresh Fruit Import demand System Parameter Estimates, 1971–1997 (Five Goods)

| Fruit | Expenditure Coefficient | Slutsky Price Coefficients | | | | | d_1 | d_2 |
		Bananas	Grapefruits	Oranges	Lemons	Others		
General system								
Bananas	.104	−.202	.157	.011	.028	.005	.300	.077
	(.192)[a]	(.072)	(.043)	(.023)	(.038)	(.014)	(.301)	(.259)
Grapefruit	.384		−.238	.058	−.002	.025		
	(.071)		(.060)	(.024)	(.023)	(.012)		
Oranges	.082			−.100	.029	.009		
	(.047)			(.033)	(.016)	(.001)		
Lemons	.076				−.062	.007		
	(.061)				(.039)	(.009)		
Others	.054					−.038		
	(.023)					(.019)		
Rotterdam system								
Bananas	.285	−.223	.168	.013	.035	.007		
	(.061)[a]	(.039)	(.037)	(.019)	(.021)	(.011)		
Grapefruit	.421		−.253	.060	.001	.024		
	(.060)		(.048)	(.023)	(.022)	(.012)		
Oranges	.100			−.107	.031	.004		
	(.044)			(.018)	(.015)	(.007)		
Lemons	.122				−.073	.007		
	(.040)				(.020)	(.008)		
Others	.072					−.041		
	(.015)					(.095)		
CBS system								
Bananas	−.322	−.212	.150	.019	.031	.011		
	(.069)[a]	(.046)	(.042)	(.020)	(.021)	(.011)		
Grapefruit	.301		−.237	.062	−.002	.028		
	(.064)		(.053)	(.023)	(.022)	(.012)		
Oranges	.042			−.111	.032	−.002		
	(.043)			(.017)	(.014)	(.007)		
Lemons	−.033				−.072	.012		
	(.039)				(.018)	(.008)		
Others	.012					−.049		
	(.016)					(.010)		

Note: Parameters are provided under the restrictions of homogeneity and symmetry.
[a] Asymptotic standard errors are in parentheses.

from zero. All own-price parameters are negative and significantly different from zero at α = .05, except that of lemons, which is statistically different from zero at α = .10. All significant cross-price terms are positive with four out of ten being different from zero at α = .05.

The estimates for the Rotterdam import demand system, shown in the middle panel of Table 4, indicate that the marginal import ex-

penditure shares are all positive and different from zero at α = .05. All own-price parameter estimates are negative, and all cross-price parameter estimates are positive. All own-price parameters are significant at α = .05, with the exception of the "others" category. Slutsky cross-price parameters are significant at α = .05 for banana-grapefruit, grapefruit-oranges, oranges-lemons, and grapefruit-others. This indicates that these four combinations of

goods are Hicksian substitutes with respect to each other.

Results of the CBS model with homogeneity and symmetry restrictions are reported in the bottom panel of Table 4. Remember that, for the CBS model, an expenditure estimate greater than, less than, or equal to zero indicates an expenditure elasticity greater than, less than, or equal to unity, respectively. The expenditure parameter for bananas is negative and significant at $\alpha = .05$, which implies that the expenditure elasticity for bananas is less than one. On the other hand, the expenditure parameter for grapefruit is positive and significant, which implies that the expenditure elasticity for grapefruit is greater than one. All other expenditure parameters for the CBS model are not significant. All own-price parameters are negative and statistically different from zero ($\alpha = .05$). All significant cross-price parameters are positive. The same combinations of goods that are substitutes in the Rotterdam model are also substitutes in the CBS model.

Expenditure Elasticity Estimates

Conditional import expenditure elasticities, conditional Slutsky price elasticities, and conditional Cournot price elasticities are provided in Table 5. The elasticities for both the Rotterdam and the CBS model are calculated from their respective parameter estimates (Table 4) with homogeneity and symmetry imposed and using the sample mean import expenditure share from 1971–1997. The asymptotic standard errors are given in parentheses. The formula for the conditional expenditure elasticity of good i associated with the Rotterdam model is $\eta_i = \theta_i/w_i$, whereas the corresponding conditional expenditure elasticity associated with the CBS model is $\eta_i = 1 + \beta_i/w_i$. The expenditure elasticity associated with the CBS model is obtained by replacing θ_i in the Rotterdam expenditure elasticity formula with $(w_i + \beta_i)$ and simplifying.

The import expenditure elasticities in Table 5 are calculated at the sample mean conditional budget shares (1971–1997) and are all statistically different from zero at $\alpha = .05$. Both

the Rotterdam and CBS estimates indicate that the conditional import expenditure elasticity for bananas and lemons is less than unity, and both indicate that the conditional import expenditure elasticity of grapefruit is greater than unity. However, under Rotterdam, the import expenditure elasticities for oranges and others are less than unity, whereas under CBS, these are greater than unity. This is good news for U.S. grapefruit exporters to Japan, because 95%–99% of Japanese grapefruit are from U.S. sources. As Japanese import expenditures for fresh fruits increase, the share of grapefruits should increase as well. However, U.S. lemon exporters will see a decline in the share of lemons imported as fruit import expenditures increase.

Own-Price Elasticity Estimates

Fruit exporters are also interested in the responsiveness of import demand to changes in the own-price of the particular type of fruit in question. Two types of own-price elasticities can be calculated from the resulting parameters: Slutsky and Cournot. Conditional Slutsky (compensated) price elasticities indicate the percentage response in quantities demanded resulting from a 1% change in price, holding real expenditures on imported fruits constant. The formula for the conditional Slutsky own-price elasticity of good i is $S_i = \pi_{ii}/w_i$. This formula is the same for both the Rotterdam and CBS models, but the empirical estimates differ because the Slutsky parameter estimates from the competing models are different.

Conditional Cournot (uncompensated) price elasticities indicate the percentage response in quantities demanded resulting from a 1% change in price, holding nominal expenditures on imported fruits constant. The formula for the conditional Cournot own-price elasticity of good i associated with the Rotterdam model is $C_i = \pi_{ii}/w_i - \theta_i$. The formula for the conditional Cournot own-price elasticity of good i associated with the CBS model is obtained by replacing the marginal import share (θ_i) with $w_i + \beta_i$ in the formula for the Cournot own-price elasticity of the Rotterdam model. This procedure results in a Cournot

Table 5. Estimated Conditional Expenditure and Price Elasticities for Rotterdam and CBS Systems Calculated at Sample Means, Five Fresh Imported Fruits, 1971–1997

Imports	Expenditure Elasticities	Slutsky Price Estimates[a]					Cournot Price Elasticities[a]				
		Bananas	Grapefruit	Oranges	Lemons	Others	Bananas	Grapefruit	Oranges	Lemons	Others
Rotterdam system											
Bananas	.61	−.48	.36	.03	.07	.02	−76	.24	−.04	−02	−04
	(.13)[b]	(.08)	(.08)	(.04)	(.04)	(.02)	(.09)	(.08)	(.04)	(.05)	(.03)
Grapefruit	2.24	.89	−1.34	.32	.00	.13	−.16	−1.77	.08	−.34	−.06
	(.32)	(.20)	(.26)	(.12)	(.12)	(.07)	(.23)	(.26)	(.13)	(.13)	(.07)
Oranges	.95	.12	.57	−1.02	.30	.03	−.32	.39	−1.12	.15	−.05
	(.42)	(.18)	(.22)	(.17)	(.15)	(.07)	(.25)	(.23)	(.19)	(.16)	(.08)
Lemons	.79	.22	.00	.20	−.47	.04	−.15	−.14	.12	−.59	−.02
	(.26)	(.14)	(.14)	(.10)	(.13)	(.05)	(.17)	(.15)	(.11)	(.14)	(.06)
Others	.85	.09	.29	.04	.08	−.49	−.31	.13	−.05	−.05	−.56
	(.18)	(.13)	(.15)	(.09)	(.10)	(.11)	(.14)	(.15)	(.09)	(.10)	(.12)
CBS system											
Bananas	.32	−.45	.32	.04	.07	.02	−.60	.26	.01	.02	.08
	(.15)	(.10)	(.09)[b]	(.04)	(.05)	(.02)	(.11)	(.10)	(.05)	(.05)	(.03)
Grapefruit	2.60	.79	−1.26	.33	−.01	.15	−.42	−1.75	.05	−.42	−.07
	(.34)	(.22)	(.28)	(.12)	(.12)	(.07)	(.26)	(.29)	(.13)	(.13)	(.07)
Oranges	1.40	.18	.59	−1.05	.30	−.02	−.47	.32	−1.20	.09	−.14
	(.41)	(.19)	(.22)	(.16)	(.13)	(.07)	(.25)	(.23)	(.17)	(.15)	(.08)
Lemons	.79	.20	−.01	.21	−.47	.07	−.17	−.16	.12	−.59	.01
	(.25)	(.14)	(.14)	(.09)	(.12)	(.05)	(.17)	(.15)	(.10)	(.13)	(.06)
Others	1.16	.13	.33	−.02	.14	−.58	−.40	.11	−.14	−.04	−.67
	(.18)	(.13)	(.15)	(.08)	(.10)	(.11)	(.14)	(.15)	(.09)	(.10)	(.12)

[a] Estimates based on parameter estimates from Rotterdam and CBS systems with homogeneity and symmetry imposed.
[b] Asymptotic standard errors are in parentheses.

own-price elasticity for the CBS model equal to $C_i = \pi_{ii}/w_i - (w_i + \beta_i)$.

Slutsky and Cournot own-price elasticities, shown in Table 5, are calculated at the sample means on the basis of parameter estimates (Table 4) from the Rotterdam and CBS models with homogeneity and symmetry (Table 5). The own-price elasticities are those along the diagonals, corresponding to the change in import quantities caused by a change in the price of the same good. The Slutsky own-price estimates from the two models are quite close in value, and all estimates are negative. The Slutsky own-price import elasticity estimates for bananas, lemons, and others are all statistically different from zero and negative, which indicates that their own-price response is inelastic. Those of grapefruits and oranges are statistically different from zero, and their point estimates are greater than unity in absolute value, which indicates an elastic own-price response. These results are important for exporters of these fruits, because they indicate whether or not an own-price change would decrease or increase revenue. For example, the own-price elasticity estimates of the Rotterdam and CBS models indicate a 1% increase in own price would decrease import demand for grapefruit 1.34% and 1.26%, respectively. The same increase in orange price would decrease demand for imported oranges by roughly 1% as indicated by both models. Accordingly, a price increase for these fruits, *ceteris paribus*, would decrease total revenue.

The own-price elasticity estimates of bananas, lemons, and others suggest the opposite. On the basis of the two models, a 1% increase in the own-price of bananas and lemons would also decrease their import demand by roughly 0.5%, whereas the same increase in the own price of other fruits would decrease import demand for others between 0.5% and 0.6%. Thus, a small increase in price would increase total revenue for bananas, lemons, and others.

The Cournot own-price elasticities provided in Table 5 are calculated by keeping nominal expenditures constant and thus are affected by price and real income effects. Accordingly, for each fruit, the Cournot esti-

mates are more negative than the corresponding Slutsky ones. However, the responsiveness of own-price changes is only slightly increased when accounting for expenditure effects of own-price changes. Point estimates for bananas, lemons, and others continue to be inelastic, whereas those of grapefruit and oranges remain elastic.

Cross-Price Elasticity Estimates

It is also important for fruit exporters to understand the effects on their product's demand from changes in price of other competing fruits. Two types of cross-price elasticities can be calculated from the resulting parameters, Slutsky and Cournot. The conditional Slutsky (compensated) cross-price elasticity of good i with respect to good j indicates the percentage response in the quantity of good i demanded resulting from a 1% change in the price of good j, holding real expenditures on imported fruits constant. The formula for the conditional Slutsky cross-price elasticity of good i with respect to good j for both the Rotterdam and CBS models is $S_{ij} = \pi_{ij}/w_i$.

The conditional Cournot (uncompensated) cross-price elasticity of good i with respect to good j indicates the percentage response in the quantity of good i demanded resulting from a 1% change in the price of good j, holding nominal expenditures on imported fruits constant. The formula for the conditional Cournot cross-price elasticity of good i with respect to good j, associated with the Rotterdam Model, is $C_{ij} = (\pi_{ij} - \theta_i w_j)/w_i$. The formula for the conditional Cournot own-price elasticity of good i with respect to good j, associated with the CBS model, is obtained by replacing the marginal import share (θ_i) with $w_i + \beta_i$ in the formula for the Cournot cross-price elasticity of the Rotterdam model. This procedure results in a Cournot cross-price elasticity of good i with respect to good j for the CBS model equal to $C_{ij} = (\pi_{ij} - \beta_i w_j)/(w_i - w_j)$.

Slutsky and Cournot cross-price elasticities calculated at sample means are also reported in Table 5. Positive Slutsky cross-price elasticities indicate that two products are substitutes, whereas negative and statistically sig-

nificant elasticities indicate complementarity. The following combinations of goods have cross-price elasticities that are statistically significant at $\alpha = .05$ under both the Rotterdam and CBS models: banana-grapefruit, oranges-grapefruit, and lemons-oranges, and grapefruit-others. All of these Slutsky cross-price elasticities are less than unity and positive, which indicates that these goods are substitutes.

The Cournot cross-price elasticity measures both price and income effects from changes in another product's price. The expenditure effect can counteract the price substitution effect, and a Cournot cross-price elasticity can be negative while the corresponding Slutsky one can be positive. The Cournot cross-price elasticities under the Rotterdam model are significant for only two combinations of goods, banana-grapefruit and others-bananas. The Cournot price elasticity of bananas with respect to grapefruit is positive, which is similar to the Slutsky price elasticity. However, the price elasticity of others with respect to bananas is negative in the Cournot case. The Cournot cross-price elasticities under the CBS model are significant for grapefruit-bananas, grapefruit-lemons, bananas-others, oranges-bananas, and oranges-others. Furthermore, some of these Cournot cross-price elasticities are positive, whereas others are negative.

Empirical Results for Six Goods

In the previous section, we presented results for Japan's four largest fresh fruit imports (bananas, grapefruits, oranges, and lemons) and other fruits, an aggregation of pineapples, berries, and grapes, into one category. We were unable to reject both the Rotterdam and CBS specification, so we presented the results from both models. In this section, we remove pineapples from the other category and reestimate the entire system by disaggregating the types of fruit into the following six categories: (1) bananas, (2) grapefruit, (3) oranges, (4) lemons, (5) pineapples, and (6) others. The "others" category now contains just berries and grapes. The purpose of this exercise is to de-

termine whether differences in the level of aggregation change the results significantly, both qualitatively and in terms of model selection. Berries and grapes remain grouped together in this section because the import expenditure shares (Table 2) are so small that elasticity estimates (which contain a constant term divided by the budget share) would be inaccurate if these were separated. The other reason is that, as the number of goods (or equations) in a demand system increases, the degrees of freedom and the power of asymptotic tests are lowered substantially.[8] For example, Laitinen showed that the probability of rejecting homogeneity, when it should not be rejected, increases as the number of goods in a system increases. The same problem also occurs for symmetry testing (Meisner).

Testing Restriction, Choice of Functional Forms, and Goodness-Of-Fit

The log-likelihood values associated with each demand system under six goods are provided in Table 6. LRTs for the different combinations of models and restrictions can be performed on the values in Table 6 in a similar fashion as in the previous section. For each of the five demand systems, homogeneity is not rejected. However, symmetry is rejected for all five models at the 95% confidence level, although, for the Rotterdam system, symmetry is not rejected at the 99% level. Recall that symmetry was not rejected for any of the five models under the five-good case at the 95% confidence level.[9] This result, rejection of

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[8] For the unrestricted system, increasing the number of total goods in the system from five to six requires estimating an additional $35–24 = 11$ parameters. Going from five to seven goods would require estimating an additional $48 − 24 = 24$ parameters. In the original version of this paper, the data set ended in 1993. Hence, degrees of freedom limitations were the other factor that led to the original estimation with only five goods.

[9] The Rotterdam system has the lowest value for the log-likelihood ratio test statistic with respect to the general system (20.6). This is higher than the critical value of 18.3 for the χ^2 with 10 degrees of freedom at the 95% level of significance but is lower than the critical value of 23.21 at the 99% level of significance.

Table 6. Log-likelihood Values and Alternatives for Japanese Fresh Fruits Imports, 1971–1997 (Six Goods)

	Models				
Restriction	General	Rotterdam	CBS	AIDS	NBR
Unrestricted	426.9 (37)[a]	424.4 (35)	422.5 (35)	424.6 (35)	425.8 (35)
Homogeneity	424.2 (32)	422.0 (30)	419.8 (30)	421.4 (30)	422.7 (30)
Symmetry[b]	411.9 (22)	411.7 (20)	404.8 (20)	384.9 (20)	388.7 (20)

[a] Number of free parameters for each model is in parentheses.
[b] Symmetry is the case in which both homogeneity and symmetry are jointly imposed.

symmetry with in the six-good case but not in the five-good case, is consistent with the results of Meisner (1979), who showed that the probability of rejecting symmetry increases as more goods are added to the system. Meisner also concluded that the power of the test for symmetry decreases as the number of goods increases. Hence, the log-likelihood test tends to reject symmetry more often than it should.

When testing for choice of functional form, we again use the log-likelihood tests performed on the general model with respect to the other four systems. The resulting log-likelihood values are reported in Table 6. On the basis of these tests, the CBS, AIDS, and NBR models are all rejected at the 95% level of significance. Hence, for the six-good case, only the Rotterdam specification is not rejected at the 95% confidence level.

The systemwide R_x^2 is also calculated for the six-good case. The results lend support to the choice of the Rotterdam system as the preferred functional form for this set of import data. The R_x^2 values for the General, Rotterdam, CBS, AIDS, and NBR are .79, .98, .83, and .98, respectively.

Parameter Estimates

Individual parameter estimates for the General and Rotterdam models in the six-good case are provided in Table 7. We did not include parameter estimates for the CBS, AIDS, or NBR models because these models were all rejected. All import expenditure coefficients are positive and significant at $\alpha = .05$ in both the general and the Rotterdam systems, with the exception of the others category. This result is consistent with the five-good case, with the

caveat that the import expenditure parameter for pineapples is also significant in the six-good case. The magnitudes of the expenditure coefficients for the general system are considerably different when comparing the five- and six-good cases. However, the expenditure coefficients for the Rotterdam system are quite similar when comparing the two pairwise.

The Slutsky own-price parameters provided in Table 7 are all negative and significant at $\alpha = .05$ for both the general system and Rotterdam model. In addition, the corresponding values for the Slutsky own-price parameters are remarkably similar when comparing the five- and six-good cases. The only exception is the others category for the Rotterdam model. In the five-good case, this parameter was not significant. However, the process of separating pineapples from berries and grapes generated a significant estimate of the own-price parameters for both pineapples and the others category.

The Slutsky cross-price parameters are also provided in Table 7. These results are generally consistent with those for the five-good case, with a few exceptions. In the five-good case (Table 4), a few cross-price parameters for the others category are significant. However, in the six-good case, none of the cross-price parameters are significant. Furthermore, when pineapples are disaggregated, the banana-pineapple cross-price parameter turns out to be significant at $\alpha = .05$ for the general model, and the grapefruit-pineapple coefficient is significant at $\alpha = .10$ for the Rotterdam model.

Elasticity Estimates

The elasticity estimates for the Rotterdam model associated with the six-good case are

Table 7. Japanese Fresh Fruit Import Demand System Parameter Estimates, 1971–1997 (Six Goods)

| Fruit | Expenditure Coefficient | Slutsky Price Coefficients | | | | | | d_1 | d_2 |
		Bananas	Grapefruit	Oranges	Lemons	Pineapple	Others		
General system									
Bananas	.312	−.207	.162	.013	.033	.267	−.011	−.073	.053
	(.148)[a]	(.043)	(.038)	(.019)	(.022)	(.043)	(.011)	(.221)	(.096)
Grapefruit	.435		−.252	.054	.003	.045	.011		
	(.067)		(.049)	(.023)	(.023)	(.030)	(.010)		
Oranges	.099			−.092	.027	.012	−.006		
	(.047)			(.022)	(.016)	(.016)	(.006)		
Lemons	.156				−.071	.003	.017		
	(.054)				(.023)	(.010)	(.013)		
Pineapples	.063					−.333	.005		
	(.025)					(.053)	(.006)		
Others	.008						−.016		
	(.036)						(.007)		
Rotterdam system									
Bananas	.272	−.218	.165	.016	.036	.009	−.008		
	(.060)	(.039)	(.037)	(.019)	(.021)	(.011)	(.009)		
Grapefruit	.430		−.256	.056	.001	.020	.013		
	(.061)		(.048)	(.023)	(.022)	(.012)	(.010)		
Oranges	.096			−.100	.030	.004	−.006		
	(.044)			(.018)	(.015)	(.007)	(.006)		
Lemons	.135				−.076	.004	.011		
	(.040)				(.020)	(.010)	(.007)		
Pineapples	.070					−.035	.004		
	(.016)					(.011)	(.006)		
Others	−.002						−.014		
	(.015)						(.007)		

Note: Parameters are provided under the restrictions of homogeneity and symmetry.

[a] Asymptotic standard errors are in parentheses.

provided in Table 8. The conditional import expenditure elasticities are provided in the first column. All of these elasticities are positive and significant, with the exception of the others category. Furthermore, the magnitudes of the elasticities that correspond to the five-good case are similar (Table 5). However, in the five-good case, the import expenditure elasticity for the others category is significant. In the six-good case, the dissagregation of pineapples resulted in an import expenditure elasticity of pineapples equal to 1.16. Hence, as the amount spent on Japanese fresh fruit imports increases, relatively more is spent on pineapples. The Slutsky price elasticities for the six-

good case are also shown in Table 8. The own-price elasticities are all negative and significant at $\alpha = .05$. Furthermore, these estimates are all similar to those for the five-good case. The Slutsky cross-price elasticities are also similar to the five-good case. Finally, both the Cournot own-price and cross-price elasticities are also similar.

Conclusions

Using annual Japanese fresh fruit import data from 1971 to 1997, this study analyzes the import patterns of Japan's seven most popular fresh fruits by implementing and testing a gen-

Table 8. Estimated Conditional Expenditure and Price Elasticities for the Rotterdam System Calculated at Sample Means, Six Fresh Imported Fruits, 1971–1997

Imports	Expenditure Elasticities	Bananas	Grapefruit	Oranges	Lemons	Pineapples	Others
Slutsky price elasticities							
Bananas	.58	−.47	.35	.03	.08	.02	−.02
	(.13)[b]	(.08)	(.08)	(.04)	(.05)	(.02)	(.02)
Grapefruit	2.29	.88	−1.36	.30	.01	.11	.07
	(.33)	(.20)	(.26)	(.12)	(.12)	(.06)	(.05)
Oranges	.91	.15	.54	−.95	.28	.03	−.06
	(.42)	(.18)	(.21)	(.17)	(.14)	(.06)	(.06)
Lemons	.87	.23	.01	.19	−.49	−.01	.07
	(.26)	(.14)	(.15)	(.10)	(.13)	(.05)	(.04)
Pineapples	1.16	.15	.34	.06	−.03	−.59	.072
	(.26)	(.18)	(.20)	(.11)	(.13)	(.18)	(.11)
Others	−.08	−.32	.54	−.25	.43	.18	−.58
	(.62)	(.37)	(.41)	(.25)	(.28)	(.26)	(.27)
Cournot price elasticities							
Bananas		−.74	.25	−.03	−.01	−.02	−.03
		(.09)	(.08)	(.04)	(.05)	(.02)	(.02)
Grapefruit		−.19	−1.79	.06	−.34	−.03	.01
		(.23)	(.27)	(.13)	(.13)	(.07)	(.05)
Oranges		−.27	.36	−1.05	.15	−.02	−.08
		(.25)	(.23)	(.18)	(.16)	(.07)	(.06)
Lemons		−.18	−.16	.10	−.63	−.06	.05
		(.17)	(.15)	(.1)	(.14)	(.05)	(.04)
Pineapples		−.39	.12	−.34	−.21	−.66	.04
		(.20)	(.20)	(.21)	(.14)	(.18)	(.10)
Others		−.28	.55	−.23	.44	.18	−.58
		(.46)	(.42)	(.27)	(.31)	(.28)	(.27)

[a] Estimates based on parameter estimates from the Rotterdam system with homogeneity and symmetry imposed.
[b] Asymptotic standard errors are in parentheses.

eral differential demand system that nested four alternative import demand specifications. When tested against the general system using the five-good case (bananas, grapefruits, oranges, and lemons and aggregating pineapples, berries, and grapes), the analysis rejects the AIDS and NBR specifications but does not reject Rotterdam and CBS models. When estimated using the six-good case (bananas, grapefruits, oranges, lemons, pineapples, and aggregating berries and grapes), the analysis rejects all specifications except the Rotterdam model.[10] Elasticity estimates are provided for

those demand specifications that the general model does not reject.

The results of the analysis have several implications for exporters of fresh fruits to Japan. It was found that if Japanese consumers were to increase their expenditure on fresh fruit imports in the future, they would spend a larger portion of their budget on the consumption of grapefruits and pineapples than they do currently. On the other hand, if Japanese consumers were to decrease their expenditure on fresh fruit imports (for example, because of a recession), they would spend a larger portion of

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[10] This result is in direct contrast to Lee, Brown, and Seale (1994), who found, when analyzing the domestic demand for aggregate commodity groups in Tai-

wan, that AIDS-type differential demand responses describe Taiwanese consumer behavior better than do other differential specifications.

their budget on bananas, oranges, and lemons. Furthermore, if the price of fresh fruit imports were to increase by a certain percentage in the future, grapefruit imports would drop by more than the percentage increase in price. Hence, lowering the price charged for grapefruit exports to Japan would increase total revenue for grapefruit exporters. Alternatively, banana, orange, lemon, and pineapple imports would drop by less than the percentage increase in price. Hence, increasing the price charged for bananas, oranges, lemons, and pineapples would increase total revenue for these exporters.

Another important result of the analysis is that Japanese consumers view certain types of fresh fruit imports as substitutes, meaning that if Good A and B are substitutes, an increase in the price of Good A would cause Japanese consumers to buy more of Good B as an alternative to Good A (all else remaining equal). It was found that oranges are substitutes for both grapefruit and lemons. It was also found that bananas and grapefruits are substitutes. These results should enable major exporters, such as citrus producers in the southern United States, to plan their pricing strategies accordingly so as to increase total revenue.

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