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THE ECONOMETRICS OF ELASTICITIES OR THE ELASTICITY OF ECONOMETRICS: AN EMPIRICAL ANALYSIS OF THE BEHAVIOR OF U.S. IMPORTS

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ABSTRACT

Fifty years of econometric modeling of U.S. import demand assumes that trade elasticities are autonomous parameters, that both cross-price effects and simultaneity biases are absent, and that expenditures on domestic and foreign goods can be studied independently of each other. To relax these assumptions, the paper assembles a simultaneous model explaining bilateral U.S. import volumes and prices. Spending behaves according to the Rotterdam model which, by design, embodies all of the properties of utility maximization and does not treat trade elasticities as autonomous parameters. Pricing behaves according to the pricing-to-market hypothesis which recognizes exporters' incentives to discriminate across export markets. Parameter estimation relies on the Full Information Maximum Likelihood (FIML) approach and uses bilateral price data for 1965-1987. According to the evidence, treating trade elasticities as autonomous parameters and ignoring the statistical implications of simultaneity and optimization impart significant biases to the structural estimates and undermine our effectiveness in addressing questions relevant to economic interactions among nations.

The Econometrics of Elasticities or the Elasticity of Econometrics: An Empirical Analysis of the Behavior of U.S. Imports

Jaime Marquez[1]

1. Introduction

Inspecting fifty years of econometric analyses of U.S. imports reveals a propensity to assume that trade elasticities are autonomous parameters, that both cross-price effects and simultaneity biases are absent, and that expenditures on domestic and foreign goods can be studied independently of each other.[2] To say that these assumptions are restrictive is elaborating on the obvious. Yet very little empirical work addresses all of these limitations despite the profession's persistent dissatisfaction with them:

> "However, there is reason to suspect that the log-linear model traditionally used in these investigations is incorrectly specified." [Murray and Ginman, 1976, p. 75]
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> "This allows us to emphasize the fact that very different behavioral models can lead to the same estimating equation for trade flows, in which case proper interpretation of parameter estimates calls for estimation of a system of equations. Unfortunately, the bulk of the empirical trade literature reports single estimating equations with only cursory reference to the theoretical structure motivating the equations." [Thursby and Thursby, 1987, p.1]

This paper eliminates these assumptions and shows the practical implications of maintaining them. To this end, I rely on a model with U.S. consumers substituting between domestic and foreign products, foreign exporters developing pricing policies suited to the U.S. market, and parameter estimates recognizing the interdependence between spending and pricing decisions. To explain

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[1] The author is a staff economist in the Division of International Finance. I have benefited from comments by William Barnett, Kenneth Clements, Russell Cooper, Andrew Dick, Neil Ericsson, Robert Feenstra, Thomas Fomby, Joe Gagnon, Reuven Glick, David Gordon, William Helkie, Dale Henderson, David Hendry, David Howard, Doug Irwin, Ulrich Kohli, Eric Leeper, Catherine Mann, Adrian Pagan, Andrew Rose, Janice Shack-Marquez, Clinton Shiells, Aris Spanos, and Alan Winters. I am also grateful to Alexander Italianer, from the Commission of the European Communities, who provided the trade data. All numerical calculations reported in this paper are obtained with the 1990 version of GREMLIN of TROLL and I am thankful to Bruce Gilsen and Intex for helping me to address technical questions. Finally, I am also grateful to Leo Kropywiansky who replicated the results of this paper in a wholly independent research effort. The views expressed in this paper are solely the responsibility of the author and should not be interpreted as reflecting those of the Board of Governors of the Federal Reserve System or other members of its staff.

[2] Appendix A surveys 51 papers written between 1941 and 1991 that study the behavior of U.S. imports. Of these papers, 71 percent rely on a logarithmic formulation, 78 percent ignore simultaneity biases or find them irrelevant, and 63 percent do not allow for cross-price effects. Overall, 43 percent of these studies maintain these three assumptions. See Orcutt (1950) for an early criticism of these assumptions; the criticisms raised by Balogh and Streeten (1951) have not received attention in the literature.

spending decisions, section 2 uses the Rotterdam model developed by Barten and Theil. This model embodies the properties associated with utility maximization and does not treat elasticities as autonomous parameters. To explain pricing decisions, the paper adopts the pricing model developed by Gagnon and Knetter which allows foreign producers to both exercise price discrimination and to update their pricing decisions in response to changes in demand. Section 3 estimates the parameters of the spending and pricing equations with the Full Information Maximum Likelihood (FIML) estimator using annual data for 1965-1987. Reliance on FIML avoids simultaneity biases, incorporates the restrictions associated with consumer demand theory, and allows pricing decisions to internalize explicitly the structure of preferences. This estimation rests on bilateral price data, a feature that recognizes the exporters' ability to vary prices according to destination. Most previous analyses rely on multilateral-price data which implicitly assume that the world consists of two economies, foreign and domestic, precluding foreign exporters from price discriminating.

To examine the implications of ignoring optimization, section 4 compares the estimated elasticities from the Rotterdam model with those of the conventional, log-linear model. According to the evidence, the elasticity differential between these two models is significant both from an economic and statistical point of view. This finding has practical implications for forecasting U.S. trade imbalances and predicting the response of U.S. imports to alternative exchange-rate developments. Finally, to emphasize the role of measurement errors, section 4 reports coefficient estimates that avoid simultaneity biases and incorporate the implications of utility maximization but use multilateral price data. According to the results, biases from measurement error are large enough to explain the historical neglect of simultaneity and optimization in empirical analysis of international trade.

2. Empirical Formulation

2.1 Spending Decisions

The analysis assumes that individuals determine their spending on foreign and domestic products by maximizing a utility function $u(q_o \ldots q_n)$ subject to $y = \Sigma\, p_j\, q_j$ where q_j represents purchases of products made in country j and p_j is the price paid in the United States for that product.[3] Differentiating the first order conditions for maximizing *any* $u(\cdot)$ and solving the associated system for quantities as a function of income and prices yields the demand equation for the ith product:

$$(1) \qquad w_{it}\, dlnq_{it} = [\partial(p_{it}\, q_{it}) / \partial y_t]\, dln(y/P)_t + \sum_{j=0}^{n} [(p_{it}\, p_{jt} / y_t)(\partial q_{it} / \partial p_{jt})]\, dlnp_{jt},$$

where

$$w_{it} = p_{it}\, q_{it} / y_t,$$
$$p_{jt} = (1+\tau_{jt})\, p_{xjt},$$
$$dP_t = \sum_{j=0}^{n} w_{jt}\, dlnp_{jt},$$

τ_j is the tariff rate (zero for purchases of U.S. products, q_n) and p_{xjt} is the dollar export price of products made in country j, $j<n$. In the absence of further restrictions, (1) cannot be rejected by the data and thus is not suitable for applied analysis. To implement (1) empirically, the Rotterdam model restricts $\mu_i = [\partial(p_{it}\, q_{it}) / \partial y_t]$ and $\pi_{ij} = [(p_{it}\, p_{jt} / y_t)(\partial q_{it} / \partial p_{jt})]$ to be invariant to changes in income and prices.[4] The term μ_i is the marginal budget share and it measures the additional amount spent on the ith good when income increases by one dollar. The term π_{ij} is the Slutsky coefficient and it measures the compensated price effect of a change in the price of the jth good on purchases of the

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[3] This formulation implicitly assumes that individuals differentiate products according to the place of production (Armington, 1969).

[4] An alternative to treating μ and π as constants is to specify the functional form of the utility function. Appendix E reports the results for the utility function generating the Almost Ideal Demand System of Deaton and Muellbauer (1980a).

ith good.[5]

Treating both μ_i and π_{ij} as autonomous transforms (1) into

$$w_{it}\, dlnq_{it} = \mu_i\, dln(y/P)_t + \sum_{j=0}^{n} \pi_{ij}\, dlnp_{jt} + r_{it}\ ,\ i=0,\dots,n, \tag{2}$$

where r_{it} is a random disturbance containing the (second-order) approximation terms induced by the assumed constancy of μ_i and π_{ij}.[6] In the absence of a universally accepted functional form for the utility function, determining the quality of this approximation is an empirical question that this paper addresses by testing whether the parameters of (2) are constant, as suggested by Byron (1984) and Barnett (1984). The income and price elasticities associated with (2) are μ_i/w_{it} and π_{ij}/w_{it}, respectively, and they vary in response to changes in expenditure shares. Thus explaining U.S. imports with constant-elasticity models amounts to assuming that either the expenditure shares are constant or the parameters of (2) are unstable, two implications that this paper examines.

For the parameters of the Rotterdam model to be consistent with utility maximization, they need to satisfy

(3) *adding-up*: $\sum_{j=0}^{n} \mu_j = 1,$

(4) *homogeneity*: $\sum_{j=0}^{n} \pi_{ij} = 0$ for $i=0,\dots,n$,

(5) *symmetry*: $\pi_{ij} = \pi_{ji}\ \forall\, i,j \ni i \neq j$, and

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[5] Critics of the Rotterdam model (Phlips 1974 and Goldberger 1987) argue that treating μ and π as invariant to changes in income and prices implies a Cobb-Douglas utility function which embodies income and price elasticities equal to one and minus one, respectively. This criticism, known as the McFadden critique (McFadden 1964), stems from not differentiating between micro and macro parameters. Indeed Barnett (1979, 1891) derives the implications for macro behavior of assuming that individuals behave according to the Rotterdam model *without* treating the associated micro parameters as invariant to prices and income. Using several theorems on stochastic limits, he derives the per-capita equations for the Rotterdam model with constant parameters. The discrepancies between the macro and micro parameters induce a second-order approximation error which has an expected value of zero. As an alternative to Barnett's approach, Mountain (1988) derives the Rotterdam model as an approximation to *any* individual's demand function by expanding an individual's optimal expenditure share around the mean of the logs of income and prices. Overall, the McFadden critique is not very relevant for empirical work with aggregate data, as Deaton and Muellbauer (1980b, p. 73) argue.

[6] Equation (2) corresponds to the *Absolute Price Rotterdam System*. For further developments and applications of this model see Barten (1964, 1965); Theil (1965); Goldberger (1969); Theil (1971), chapters 7 and 11; Deaton (1974); Clements and Theil (1978); and Theil and Clements (1978, 1987).

(6) *quasiconcavity*: $[\pi_{ij}]$ is negative semidefinite with rank n.

In addition to testing these restrictions, the analysis also tests Preference Independence.[7] According to this hypothesis, price effects are proportional to marginal budget shares:

$$\text{(7)} \qquad \begin{aligned} \pi_{ij} &= -\phi\, \mu_i \mu_j \\ \pi_{ii} &= \phi\, \mu_i (1-\mu_i) \end{aligned}$$

$\forall\, i,j \ni i \neq j$ where $\phi<0$ is the reciprocal of the income elasticity of the marginal utility of income. From a theoretical standpoint, (7) constrains cross-price effects to be positive and rules out complementarities among products. Thus (7) is suitable for analyses explaining the behavior of broad aggregates where complementarities are unlikely to arise. From an econometric standpoint, substituting (7) into (2) reduces the number of parameters but demands non-linear FIML for parameter estimation.[8]

Insisting on utility maximization as an organizing device has two practical advantages. First, it ensures that the demand equations for foreign and domestic products are interrelated. Most previous empirical formulations of U.S. imports do not recognize that specifying the demand for foreign products carries implications for explaining the demand for domestic goods and that neglecting these implications affects all the coefficient estimates (see Winters, 1984). Second, it recognizes that income and price elasticities are not invariant to reallocations of expenditures such as those generated by the implementation of discriminatory tariffs. Thus, unless preferences are Cobb-Douglas, constant-elasticity models are not suited to predicting the effects of changes in trade policies.

The spending decisions developed here suffer, however, from several limitations. First, decisions regarding labor supply and asset holdings are treated as separable from decisions to purchase

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[7] Preference Independence is another name for Direct Additivity, a specification of the utility function introduced by Houthakker (1960). To implement this specification empirically, the paper follows Theil and Clements, 1987, chapter 3, section 3.20.

[8] Note that imposing (7) is not equivalent to guaranteeing that the estimated own-price effects are negative or that the regularity conditions of the utility function are met. Indeed, ϕ is freely estimated and thus finding that $\hat{\phi}>0$ constitutes a rejection of the Rotterdam model by the data.

foreign and domestic products. Second, the structure of preferences ignores intertemporal substitution. These two limitations are important but they apply with the same force to the existing literature, which does not enjoy the advantages offered here. Third, spending decisions ignore the product composition of U.S. imports. Eliminating this limitation involves increasing the number of equations to allow tradeoffs across countries and products but, given the annual frequency of the data, such an extension demands more observations than are currently available.

2.2 Pricing Decisions

To explain the price of domestic products, p_{nt}, the paper assumes that domestic firms have a Cobb-Douglas production function with capital and labor as the sole productive factors. Assuming that domestic firms minimize their costs yields their pricing rule as

$$ln(p_{nt} / r_{nt}) = \theta_n + \gamma_n \, ln(w_{nt} / r_{nt}) + e_{nt} ,$$

where r_{nt} is the rental rate of the capital stock, w_{nt} is the wage rate, and e_{nt} is a random disturbance.

The analysis explains U.S. import prices with the pricing-to-market model developed by Gagnon and Knetter (1990). According to their analysis, foreign exporters discriminate across destinations and set their prices according to

$$lnp_{xit} = \theta_i + \beta_i \, lnC_{it} + \delta_i \, lnE_{it} + \gamma_i \, lnp_{nt} + e_{it}, \tag{8}$$

where C_{it} is the (foreign currency) marginal cost of firms in the ith country, E_{it} is the nominal exchange rate for the ith currency (foreign currency/\$), and e_{it} is a random disturbance. According to (8), foreign exporters increase p_{xit} in response to increases in marginal costs (β_i>0) and appreciation of their currencies (δ_i<0) where δ_i is known as the "passthrough coefficient." Finally, (8) assumes that foreign firms price their exports according to the characteristics of the U.S. domestic market as summarized by the price of that market, p_{nt}. Thus I interpret γ_i as the degree of pricing-to-market.

As formulated, (8) assumes that foreign producers insist on passing a fixed percentage δ_i of changes in exchange rates to their export prices even if the demand for their product becomes price elastic. To recognize that the exchange-rate passthrough is sensitive to changes in the price elasticity of demand, the analysis follows Gagnon and Knetter (1990, p.6) and finds that (see Appendix B)

(9) $$\delta_{it} = \{[\pi_{ii}/(2\pi_{ii} - \mu_i w_{it})] - 1\},$$

where $w_{it} = p_{xit}(1+\tau_{it})\, q_{it} / y_t$. In addition to internalizing the structure of preferences into pricing decisions, (9) has several properties of interest. First, δ_{it} responds to changes in export prices from alternative suppliers because q_{it}, and thus w_{it}, depends on *all* prices.[9] Second, the degree of exchange-rate passthrough is directly related to the associated expenditure share with δ_{it} approaching $\{[\pi_{ii}/(2\pi_{ii} - \mu_i)] - 1\}$ as w_{it} approaches one. Thus even a monopolist faces a limit on its ability to pass changes in the exchange rate to p_{xit}. Third, if income effects are small ($\mu_i \to 0$), then $\delta_{it} = -1/2$. Fourth, if price effects are absent ($\pi_{ii} \to 0$), then $\delta_{it} = -1$--that is, sellers stabilize prices in their currencies and fully pass changes in exchange rates to their dollar export prices.

From an empirical standpoint, the validity of (9) can be tested by substituting it into (8) and evaluating the change in the value of the likelihood function for the whole spending-pricing system. Performing this substitution gives

(10) $$lnp_{xit} = \theta_i + \beta_i lnC_{it} + \{[\pi_{ii}/(2\pi_{ii} - \mu_i w_{it})] - 1\} lnE_{it} + \gamma_i lnp_{nt} + e_{it}.$$

According to (10), export prices depend on the level of sales, as reflected in the expenditure share, and this dependence makes the system truly simultaneous: Prices influence quantities and vice versa. Previous empirical work on pricing to market recognizes the importance of the structure of preferences but does not internalize it for parameter estimation, a limitation that this paper avoids.[10]

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9 Changes in U.S. economic activity also induce changes in expenditure shares with a corresponding effect on the passthrough coefficient.

10 See Baldwin (1988), Knetter (1989), Marston (1990), and Gagnon and Knetter (1990). Reliance on limited information estimation methods might eliminate the simultaneity biases of these studies, but cannot internalize explicitly the sensitivity of the passthrough coefficient to the specification of preferences. To judge the importance of taking into account explicitly the structure of preferences, appendix E reports the coefficient estimates for the price equations using the Almost Ideal Demand system as an alternative representation of preferences.

3. Econometric Estimation

3.1 Data

The differentiating feature of this study's data set is the availability of prices for bilateral trade. These prices recognize the ability of foreign firms to discriminate across destinations and facilitate the estimation of cross-price effects of individuals' demand functions. In contrast, multilateral prices assume that exporters face only one export market, an assumption that handicaps the estimation of these cross-price effects. To facilitate parameter estimation, the analysis focuses on per-capita purchases of both domestic products and imports from Canada, Japan, Germany, and the Rest of the World. Purchases of domestic products are measured as real GNP minus exports in real terms; purchases of foreign goods are measured as the value of bilateral imports deflated by the corresponding bilateral price. The data for bilateral prices and the associated trade flows are annual for 1965-1987, denominated in U.S. dollars, and come from the Commission of the European Communities. Precise definitions and data sources appear in appendix C.

To emphasize the increased penetration of foreign products in the U.S. market, table 1 shows their share of U.S. expenditures rising from 3 percent in 1965 to 8 percent in 1987. Table 1 also shows that competition among foreigners has been substantial with Japan's share in the U.S. import market increasing at the expense of all other countries. Central to this competition is the ability of foreign firms to develop pricing strategies suited to the U.S. market. Indeed, if the U.S. market were just another market, then bilateral and multilateral prices would have the same informational content with their changes being identical in direction and similar in magnitude. Yet figure 1 reveals that these two conditions are not met. For example, Japan's bilateral export price to the United States declines by 18 percent in 1976 whereas Japan's multilateral export price increases by 3 percent. Similarly, Germany's bilateral export price to the United States is virtually constant during 1981-1983 whereas Germany's multilateral export price exhibits declines during this period as large as 18 percent. These differences in price behavior motivate this study's reliance on bilateral trade prices for parameter estimation.

Table 1

U.S. Expenditure and Import Shares[1]

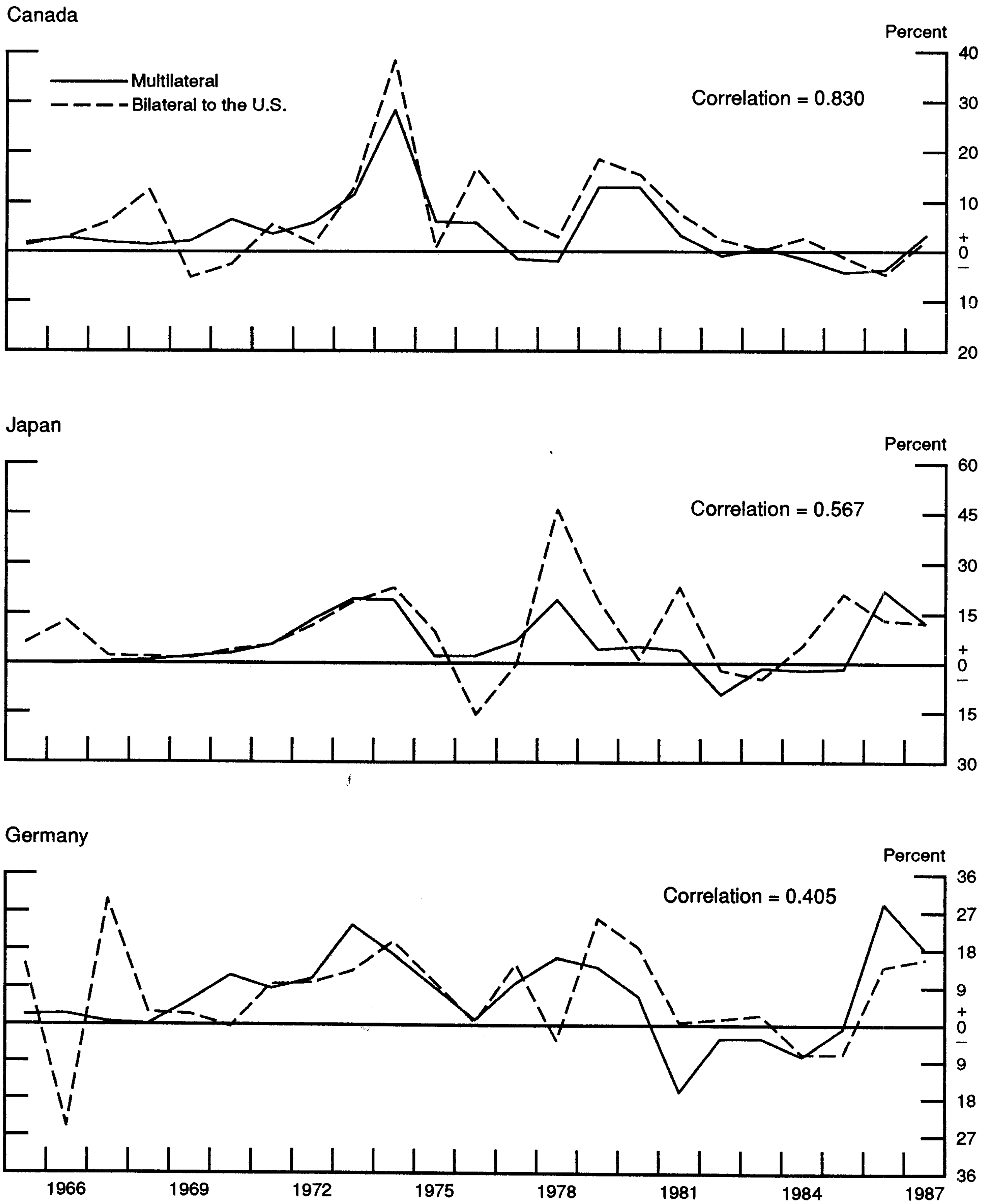
	Expenditure Shares					Import Shares				
	Canada	Japan	Germany	R.O.W.	Domestic	Canada	Japan	Germany	R.O.W.	Growth in Real Expenditures[2]
1965	0.76	0.38	0.21	1.63	97.02	25.53	12.76	7.09	54.62	6.07
1966	0.88	0.42	0.26	1.73	96.71	26.62	12.89	7.81	52.68	6.08
1967	0.96	0.40	0.26	1.69	96.69	28.93	12.22	7.97	50.87	2.79
1968	1.09	0.50	0.33	1.85	96.23	29.01	13.19	8.84	48.96	4.37
1969	1.18	0.55	0.30	1.78	96.19	30.93	14.55	7.75	46.77	2.49
1970	1.20	0.63	0.34	1.80	96.04	30.16	15.98	8.50	45.36	-0.68
1971	1.26	0.72	0.36	1.82	95.83	30.27	17.31	8.71	43.72	3.26
1972	1.33	0.81	0.38	1.91	95.58	29.99	18.24	8.55	43.22	4.88
1973	1.40	0.77	0.43	2.35	95.05	28.25	15.63	8.63	47.49	3.86
1974	1.64	0.92	0.47	2.93	94.04	27.54	15.39	7.94	49.12	-1.65
1975	1.55	0.80	0.38	2.61	94.67	29.04	14.97	7.09	48.91	-1.55
1976	1.63	0.96	0.35	2.79	94.28	28.55	16.69	6.07	48.70	5.51
1977	1.61	1.02	0.40	3.00	93.97	26.69	16.95	6.60	49.76	5.12
1978	1.64	1.27	0.51	3.52	93.07	23.68	18.25	7.29	50.77	5.27
1979	1.67	1.22	0.50	3.85	92.76	23.02	16.84	6.95	53.20	0.65
1980	1.63	1.31	0.49	4.13	92.44	21.54	17.35	6.53	54.58	-1.42
1981	1.63	1.43	0.43	3.79	92.73	22.44	19.64	5.87	52.05	1.85
1982	1.57	1.37	0.43	3.60	93.03	22.50	19.65	6.15	51.69	-1.52
1983	1.62	1.37	0.42	3.54	93.05	23.38	19.73	5.99	50.90	4.98
1984	1.84	1.70	0.50	3.91	92.05	23.12	21.36	6.30	49.22	7.80
1985	1.78	1.89	0.56	3.76	92.02	22.25	23.73	6.96	47.07	3.78
1986	1.65	2.12	0.65	3.66	91.92	20.38	26.30	8.01	45.32	2.28
1987	1.65	2.06	0.65	3.90	91.74	19.98	24.91	7.93	47.18	2.22

Sources: Data for International Trade flows are from the Volimex Data Tape provided by the Commission of the European Communities. Data for U.S. domestic variables come from the Bureau of Economic Analysis. R.O.W. stands for the Rest of the World which, in the present study, excludes OPEC and Centrally Planned Economies.

[1] These shares measure total expenditures as the market value of domestic spending on domestic goods plus the value of imports inclusive of tariffs. This valuation of imports ensures that domestic production and imports reflect market prices. Shares may not total 100 because of rounding.

[2] Growth in real expenditures is constructed as the annual growth rate of (y/P); see equation (1).

Figure 1
Annual Growth Rates of Export Prices
(U.S. dollars)



3.2 Empirical Results

Assuming that the disturbances of the spending and pricing equations have a joint normal distribution with zero mean and constant covariance matrix, the analysis estimates the associated parameters with FIML using annual observations for 1965-1987. To avoid the singularity that (4) induces, the estimation drops the equation for purchases from the rest of the world, q_{ot}.[11] As a whole, the equations of the estimating system explain U.S. per-capita purchases and prices of four expenditure categories: domestic products and imports from Canada, Germany, and Japan. The exogenous variables are nominal income, four nominal wage rates, three nominal exchange rates, the tariff rate, the rental rate for U.S. capital, and the price of imports from the rest of the world, p_{ot}, which is the numeraire.[12]

Table 2 presents the log-likelihood ratio tests associated with the hypotheses of homogeneity, symmetry, preference independence, and profit maximization;[13] I test parameter constancy for each of these hypotheses.[14] Given the sequential nature of these tests, I adopt an overall significance level of five percent which gives approximately a one percent significance level for sequential pairwise tests. Based on the evidence, the data cannot reject the restrictions associated with homogeneity and

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[11] The associated parameter estimates are recovered using equations (3)-(5).

[12] For parameter estimation the analysis replaces the differentials with differences. To avoid the associated ambiguities in the choice of initial conditions, the analysis follows Theil (1971, p. 331) and uses $w^*_{it} = (w_{it} + w_{i,t-1})/2$ instead of w_{it} in (2). This substitution introduces an additional approximation error into the disturbance of (2). Note that the estimation treats expenditure shares as endogenous variables. The estimation algorithm minimizes the negative of the concentrated log-likelihood function following the procedure of Broyden, Fletcher, Goldfarb, and Shanno (BFGS); see Hendry, Neale, and Srba (1988) for a description.

[13] The adding-up restriction, equation (3), holds by construction and thus is not subject to empirical testing. To determine whether quasiconcavity in $u(\cdot)$ holds, equation (6), the analysis computes the eigenvalues of the Slutsky matrix and tests whether they are positive. If this hypothesis is rejected, then the data cannot reject quasiconcavity in $u(\cdot)$.

[14] To test parameter constancy, the analysis tests whether the expected forecast error for 1986-1987 is zero. To this end, each equation is expanded to include two dummy variables, one for each of these two years. Testing the hypothesis that the expected forecast error is zero over this period amounts to testing that the coefficients of all of these dummy variables (2x8=16 in total) are jointly equal to zero. Note that the dollar depreciated sharply over these two years. Thus if hysteresis is important empirically, then the expected value of the forecast errors for 1986-87 should be non-zero. Moreover, substantial changes in the explanatory variables, such as this dollar depreciation, offer an opportunity to evaluate the quality of the Rotterdam model as an approximation. Again, a poor approximation will yield non-zero forecast errors.

Table 2
Preference Structure, Profit Maximization, and Parameter Constancy:
Log-likelihood ratio tests

Hypothesis[1]	No. of Parameters	Log-likelihood $L(\Phi_i)$	Test Statistic[2]	χ^2 Critical Value[3]
1. No Constraints	38	12.30		
1a. Parameter Constancy	54	27.40	30.20	32.0
2. Homogeneity (equation 4)	34	7.48	9.64	13.3
2a. Parameter Constancy	50	23.14	31.32	32.0
3. Symmetry (equation 5)	28	1.29	12.38	16.8
3a. Parameter Constancy	44	14.28	25.98	32.0
4. Preference Independence (equation 7)	19	-9.34	21.26	21.7
4a. Parameter Constancy	35	5.75	30.18	32.0
5. Profit Maximization (equation 9)	16	-12.77	6.86	11.3
5a. Parameter Constancy	32	2.42	29.86	32.0

[1] The restrictions for the ith hypothesis include the restrictions embodied in the (i-1)th hypothesis for i>1.

[2] The test statistic is $-2[L(\hat{\Phi}_i) - L(\hat{\Phi}_{i-1})]$ where $L(\hat{\Phi}_i)$ is the value of the concentrated log-likelihood function for the ith hypothesis, and $\hat{\Phi}_j$ is the vector of parameter estimates associated with the jth hypothesis (j=2,...,5). For parameter constancy, the test-statistic is $-2[L(\hat{\Phi}_j) - L(\hat{\Phi}_{ja})]$.

[3] The critical values correspond to the 1 percent significance level for a χ^2 with degrees of freedom equal to the number of additional restrictions.

symmetry. In addition, the data cannot reject preference independence but the gap between the test-statistic and the associated critical value is small. However, given the frequency with which the restrictions from consumer demand theory are rejected empirically,[15] I interpret this failure to reject as providing reasonable support for the associated restrictions. Given preference independence, the data cannot reject the implications of profit maximization as reflected in equation (9). Finally, the data cannot reject parameter constancy for each of these parametric configurations. This finding suggests that the (second-order) approximation errors in the disturbances of (2) are sufficiently small to treat the Rotterdam model as an adequate approximation to the U.S. per-capita demand functions.

The parameter estimates of the model embodying both preference independence and profit maximization exhibit several features of interest (table 3). First, marginal budget shares are positive, significant, and vary from 0.009 for imports from Germany to 0.89 for purchases of U.S. products revealing a bias in favor of domestic products. Second, own-price effects are negative, significant, and vary from -0.003 for purchases from Germany to -0.035 for purchases of U.S. products. Third, substitutability *among* foreign products is quantitatively small but statistically significant. Substitution *between* foreign and domestic products is, however, large and statistically significant.[16] Fourth, the eigenvalues of the Slutsky matrix are negative, as implied by the second order conditions for utility maximization. Fifth, $\hat{\phi}$ is significantly less than zero as implied by consumer demand theory. Sixth, changes in U.S. prices lead to proportional changes in the export prices of Canada and Germany and more than proportional changes in Japan's export price. Seventh, 85 percent of the value added in the production of domestic goods is earned by labor and 15 percent by capital.

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[15] See the discussion by Phlips (1974), Laitinen (1978), Deaton and Muellbauer (1980b), and Theil and Clements (1987).

[16] The relatively small magnitude of the cross-price effects is due to the assumption of preference independence. Relaxing this assumption raises the magnitude of these estimates but their standard errors increase more than proportionally because of the relatively large number of parameters being estimated. These results do not settle the controversy of whether cross-price effects are important in explaining the behavior of U.S. imports. For example, Stone (1979) does not find these cross-price effects to be important whereas Rousslang and Parker (1984), using non-parametric estimation methods, reach the opposite conclusion. For a further discussion of this issue, see Rousslang (1989) and Shiells, Stern, and Deardorff (1989).

Table 3

Coefficient Estimates for Spending and Pricing Decisions: FIML, 1965-1987[1]

	Spending Decisions						Pricing Decisions	
	Marginal Budget Share μ_i	Slutsky Matrix $[\pi_{ij}]$					U.S. Price γ_i	Labor Costs β_i
		Canada	Japan	Germany	U.S.	Eigenvalues[2]		
Canada	0.032878	-0.011908	0.000277	0.000116	0.011019	-0.041005	1.007140	0.404190
	0.005329	0.001647	0.000064	0.000025	0.001588	0.001174	0.506029	0.476650
Japan	0.022467		-0.008225	0.000079	0.007530	-0.010608	1.420310	-0.054428
	0.006579		0.001424	0.000028	0.001383	0.001593	0.093044	0.130876
Germany	0.009397			-0.003486	0.003149	-0.004625	1.074730	-0.047709
	0.002654			0.000717	0.000645	0.002316	0.228637	0.320423
U.S.	0.894879				-0.035230	-0.002020		0.849215
	0.024484				0.004337	0.005139		0.023214
$\hat{\phi}$	-0.374511							
	0.087408							
Average Expenditure Share		0.0144	0.0107	0.0042	0.9422			

[1] For a given cell, the top entry is the FIML estimate and the bottom entry is the associated standard error. See appendix D for the computation of the standard errors; ϕ is the reciprocal of the income elasticity of the marginal utility of income.

[2] See appendix D for the computation of the standard errors.

Finally, the estimated passthrough coefficients (figure 2) are statistically less than -1/2, a finding that reflects the steady penetration of foreign products into the U.S. domestic market.[17] Moreover, these coefficients do not change significantly after 1984, a finding contradicting the view that foreign producers reacted to the depreciation of the dollar by lowering the degree of exchange-rate passthrough to maintain their market shares.[18]

4. Sensitivity Analysis

4.1 The Econometrics of Elasticities

What are the trade elasticities implied by the estimates of the Rotterdam model? How much do they change over time? Based on their 95 percent confidence intervals, income elasticities from the Rotterdam model are positive, significant, and decline over time with a marked tendency towards homotheticity (figure 3). Specifically, income elasticities decline from 5.1 to 2.0 for imports from Canada, from 6.0 to 1.0 for imports from Japan, and from 4.5 to 1.3 for imports from Germany. These declines confirm earlier findings of parameter instability in constant-elasticity models. For example:

> "Previous investigations of imports had suggested evidence of structural change in the mid-1960s on the basis of split samples but without an unambiguous dating as to when the change may have occurred. We used a series of tests that permitted the data to indicate the presence of structural change. Our results give some support to the earlier findings that change occurred in the mid- to late 1960s." [Stern, Baum, and Green, 1979, p. 191]

Indeed the declines in income elasticities of figure 3 are most pronounced during the 1960s. But unlike previous studies, the changes in elasticities shown in figure 3 stem from the increased market penetration of foreign products and not from parameter instability in either the spending or the pric-

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[17] The point estimate of the ith passthrough coefficient is $\{[\hat{\phi}(1 - \hat{\mu}_i)/(2\hat{\phi}(1 - \hat{\mu}_i) - w_{it})] - 1\}$. Appendix D explains the construction of the associated standard errors.

[18] See the review by Hooper and Mann (1989).

Figure 2
Passthrough Coefficients
95% Confidence Intervals

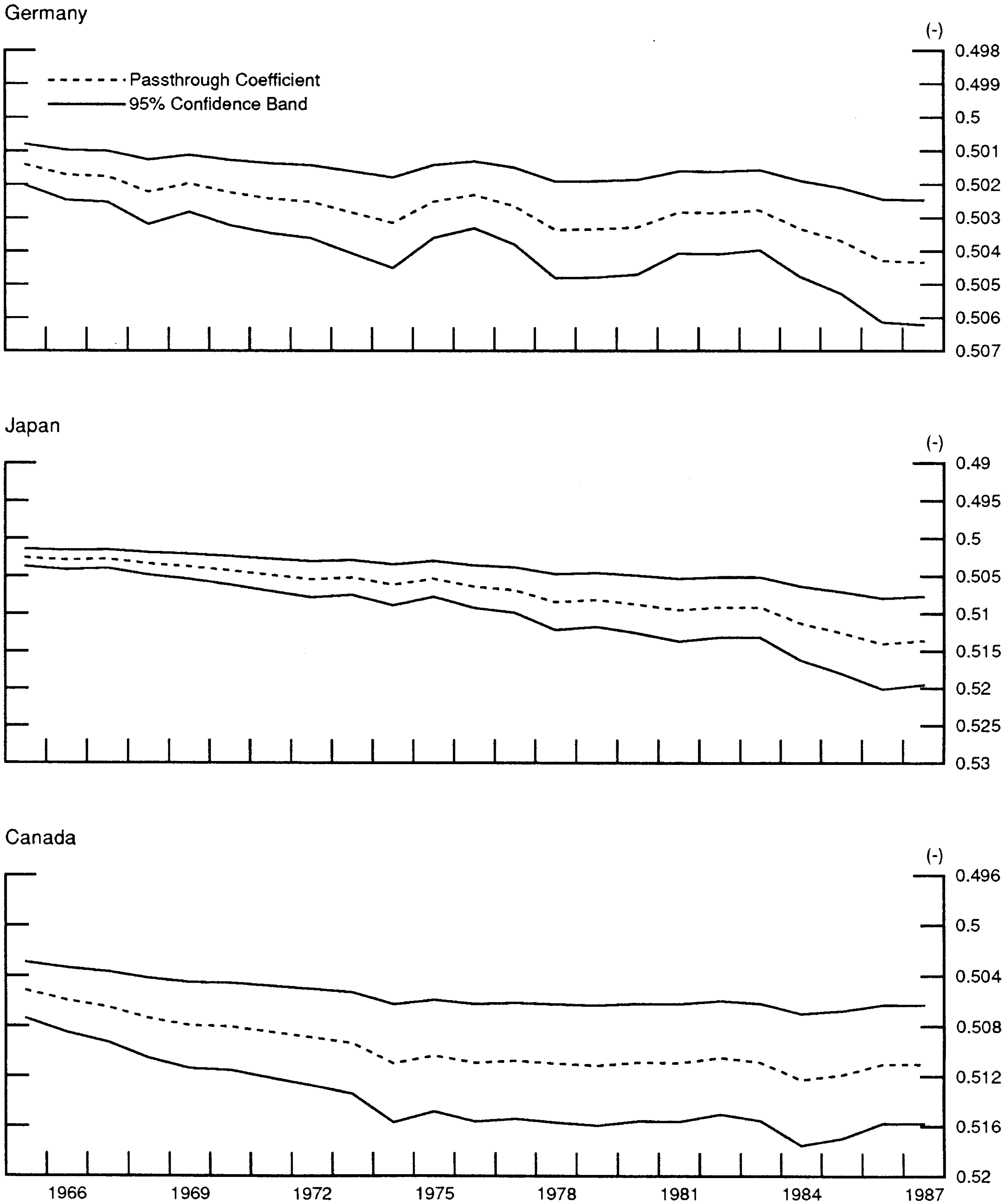
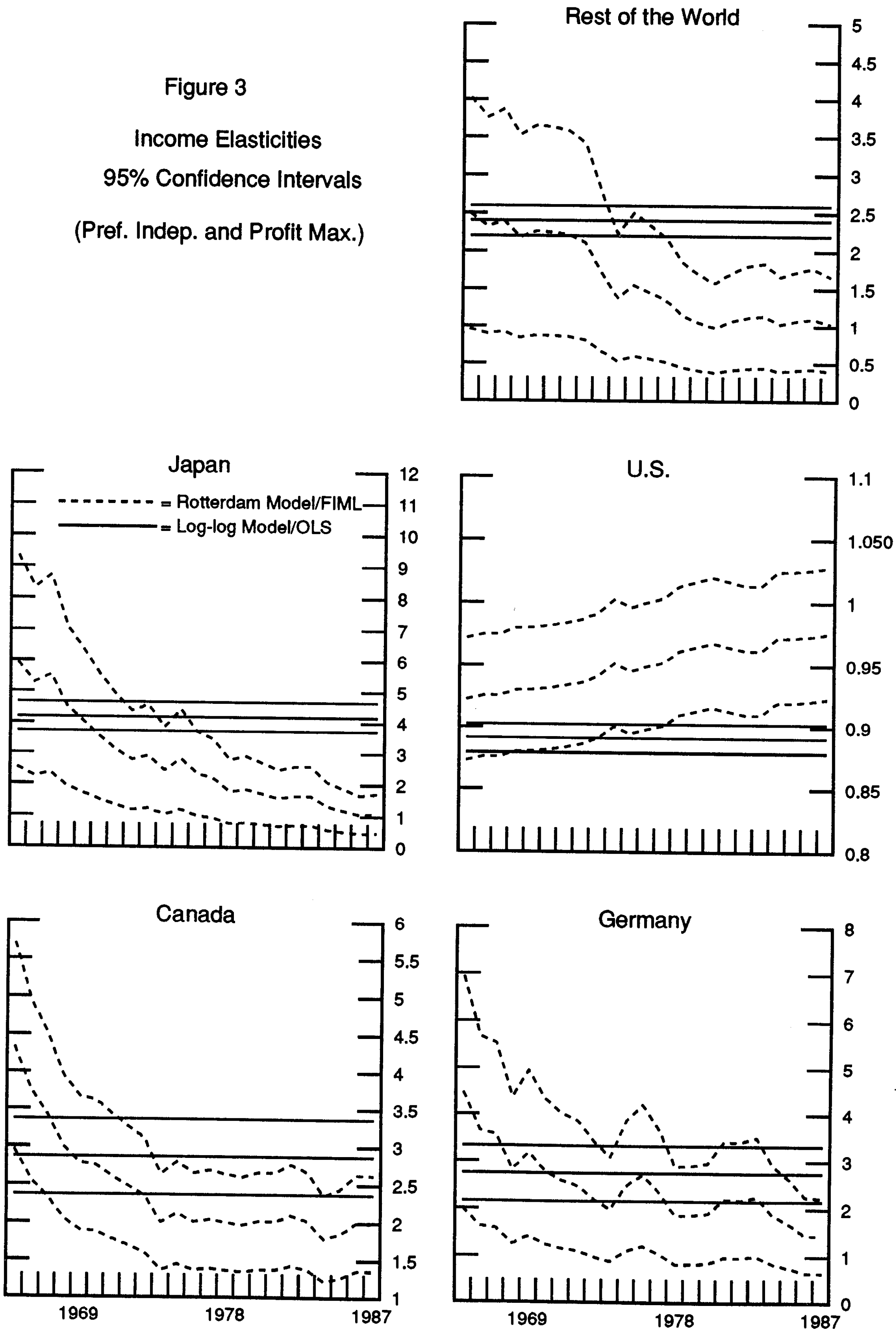


Figure 3

Income Elasticities

95% Confidence Intervals

(Pref. Indep. and Profit Max.)



ing equations.[19]

Similarly, the compensated own-price elasticities from the Rotterdam model are negative, significant, and increase from -1.6 to -0.7 for imports from Canada, from -2.3 to -0.4 for imports from Japan, and -1.6 to -0.5 for imports from Germany (figure 4). These increases confirm the delinking of U.S. imports from prices that Krugman (1989) notes. The evidence, however, does not support Krugman's hypothesis that this delinking stems from the volatility of nominal exchange rates. If this volatility were responsible for the delinking of U.S. imports, then price elasticities should stay relatively constant during the Bretton-Woods period and fall immediately afterwards whereas the results from the Rotterdam model show that price elasticities decline *during* the Bretton-Woods period. The evidence also reveals that the price elasticity for U.S. products is substantially smaller than the corresponding import-price elasticities. This smallness is one instance where reliance on elasticities, as opposed to propensities, is a nuisance. Specifically, purchases of U.S. products have the largest Slutsky coefficient but the smallest price elasticity because U.S. purchases comprise the lion's share of expenditures. Thus large price effects are not equivalent to large price elasticities.

Overall, the evolution of these elasticities suggests that when an economy is virtually closed to international trade, as the United States during the 1960s, purchases of foreign products behave as purchases of luxuries in the sense of encountering an abundance of domestic substitutes, and thus exhibiting high income and price elasticities. But as purchases of foreign products become more important in domestic transactions, imports behave more as necessities than as luxuries with a corresponding decline in trade elasticities. Note, however, that the Rotterdam model does not impose any restriction on the behavior of expenditure shares (other than adding up to one) and therefore the associated elasticities are not constrained to follow any particular path. Trade elasticities could change in response to a realignment of expenditures induced by changes in tariffs, factor prices, and exchange rates.

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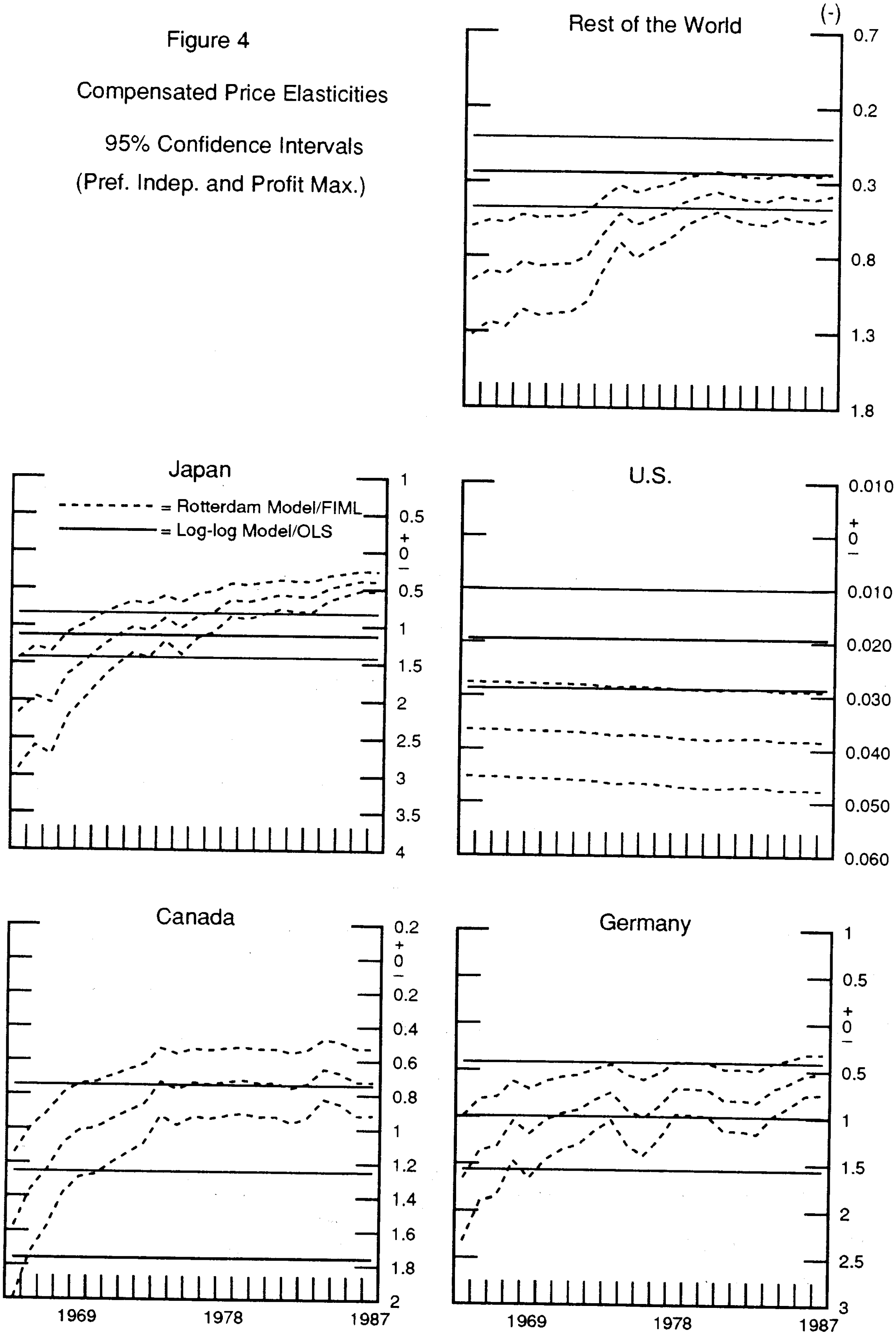
[19] See Volker (1982) for a critique of the evidence reported by Stern et al. Joy and Stolen also report instability in income elasticities; see Hooper (1977) for a comment on their paper.

Figure 4

Compensated Price Elasticities

95% Confidence Intervals

(Pref. Indep. and Profit Max.)



How do these elasticity estimates compare with those of the log-linear model? To address this question, figures 3 and 4 also report the 95 percent confidence intervals for the OLS estimates of

$$(11) \qquad lnq_{it} = \eta_i \, ln(y / P)_t + \Sigma_j \, \varepsilon_{ij} \, lnp_{jt},$$

where $\Sigma_j \varepsilon_{ij} = 0$ for $i, j = 0, \ldots n$. Equation (11) shares with the literature (see appendix A) the choice of both functional form and estimation method but it differs from previous analyses in two respects: the presence of cross-price effects to avoid misspecification biases and the use of domestic absorption in real terms instead of real GNP.[20] Note that for (11) to be consistent with utility maximization, η_i and ε_{ij} must equal one and minus one, respectively (Deaton and Muellbauer, 1980b, p. 17).

According to the results, the gap between the estimates from (11) and those from the Rotterdam model grows over time and is significant for most of the expenditure categories. For example, the income elasticity for U.S. imports from Japan generated by (11) is 4.25 whereas the income elasticity generated by the Rotterdam model for 1987 is 1.0; similarly, the own-price elasticity for imports from Japan ranges from -0.4 for the Rotterdam model to -1.1 for (11). These differences in elasticities have important implications for understanding U.S.-Japan trade relations. For example, if the question is how sensitive are U.S. imports from Japan to changes in U.S. economic activity, then the estimates from (11) greatly overstate the response of these imports relative to the predictions from the Rotterdam model. Similarly, relying on the estimated price elasticities from (11) exaggerates the degree to which Japanese products can be substituted by U.S. products. These observations would apply with the same force to imports from Canada and Germany but the standard errors associated with the own-price elasticities from (11) are sufficiently large to mask significant differences. Finally, the income elasticities from (11) are significantly different

[20] Using real GNP in (11), as opposed to real absorption (y / P), implicitly assumes that the share of imports in expenditures is sufficiently small to justify ignoring the income effects of changes in import prices. For example, equation (11) predicts that a change in p_j has a substitution effect, as measured by ε_{ij}, and an income effect, as measured by $-\eta_i w_j$. Using real GNP amounts to treating w_j as zero, an increasingly unrealistic assumption for the United States as table 1 indicates.

from one, suggesting that the evidence does not support the view that (11) is consistent with utility maximization. Overall, the gap in elasticity predictions between the log-linear model and the Rotterdam model is large enough to argue against explaining U.S. imports with constant-elasticity models.

4.2 The Elasticity of Econometrics

What do we lose by studying expenditures on foreign and domestic products as though they were independent of each other? How large are the simultaneity biases? How important is the use of bilateral prices for parameter estimation? To address the first question, I estimate (2) with and without the interdependencies implied by consumer demand theory as reflected in equation (5). To ensure that differences in these estimates are not due to the modeling of price behavior, this comparison treats prices as exogenous variables. According to the results (table 4), ignoring these interdependencies lowers the own-price effects for purchases of Canadian and U.S. products by 19 percent and more than 300 percent, respectively (compare lines 7 and 8); marginal budget shares are, however, robust (lines 1 and 2).

To examine the implications of ignoring simultaneity, the analysis compares FIML estimates that abstract from price determination with FIML estimates that internalize price behavior. The results reveal that, as long as the implications of optimization are taken into account, neglecting simultaneity leads to small biases for the parameter estimates of the Rotterdam model (compare lines 2 and 3, and lines 8 and 9). The two exceptions to this pattern are the 25 percent reduction in the own-price effect for imports from Canada (lines 8 and 9) and the 27 percent increase in the marginal budget share for imports from Germany (lines 1 and 2). Neglecting simultaneity, however, reduces the pricing-to-market coefficient for Canada to virtually zero (compare lines 13 and 14). Given that a large share of U.S. imports come from Canada (table 1), characterizing Canadian exporters' pricing policy to the U.S. market as being independent of changes in U.S. prices is likely to be an important source of error when predicting how the external imbalance of the United States would respond to changes in U.S. prices.

Table 4

Sensitivity of Estimates to Price Data and Estimation Method[1]

	Canada	Japan	Germany	U.S.
Marginal Budget Share (μ_i x 100)				
Bilateral Prices				
1. OLS and Homogeneity	3.22	3.09	1.28	84.40
2. FIML and Prices Exogenous	3.14	3.32	1.19	86.71
3. FIML and Prices Endogenous	3.29	2.25	0.94	89.49
Multilateral Prices				
4. OLS and Homogeneity	3.80	3.21	1.15	76.41
5. FIML and Prices Exogenous	3.69	3.72	1.28	75.12
6. FIML and Prices Endogenous	3.94	4.10	1.58	75.51
Compensated Own-Price Effect (π_{ii} x 100)				
Bilateral Prices				
7. OLS and Homogeneity	-0.75	-0.87	-0.34	-0.70*
8. FIML and Prices Exogenous	-0.89	-0.94	-0.34	-3.36
9. FIML and Prices Endogenous	-1.19	-0.82	-0.35	-3.52
Multilateral Prices				
10. OLS and Homogeneity	-0.20*	-0.59*	-0.36	-2.04*
11. FIML and Prices Exogenous	-0.46	-0.46	-0.16	-2.40
12. FIML and Prices Endogenous	-0.002	-0.002	-0.001	-0.01
Pricing-to-market (γ_i)				
Bilateral Prices				
13. OLS	0.0456*	1.3139	1.1225	
14. FIML and Prices Endogenous	1.0071	1.4203	1.0747	
Multilateral Prices				
15. OLS	0.1577*	0.2095	0.2431	
16. FIML and Prices Endogenous	0.6547	0.2075	0.3147	

Memorandum: Estimates of $\hat{\phi}$	$\hat{\phi}$
Bilateral Prices	
FIML and Prices Exogenous	-0.2920
FIML and Prices Endogenous	-0.3745
Multilateral Prices	
FIML and Prices Exogenous	-0.1282
FIML and Prices Endogenous	-0.0005

[1] The estimation sample is 1965-1987. An asterisk denotes that the estimate is not significantly different from zero at the five percent significance level.

Finally, to emphasize the advantages of bilateral price data, table 4 reports estimates based on multilateral prices.[21] According to the results, reliance on these prices yields a spectacular deterioration of the estimates even if they avoid simultaneity biases and incorporate the implications of utility maximization. For example, the own-price effects estimated with multilateral-price data are negligible which contrasts with the estimates based on bilateral-price data. Similarly, the marginal budget share for purchases of domestic products drops from 0.89 using bilateral-price data to 0.76 using multilateral-price data. This difference in estimates stems from an internal inconsistency between data construction and model design: multilateral prices assume that exporters do not price-to-market and that cross-price effects are zero whereas the model of this paper relaxes these two assumptions. This inconsistency affects all of the estimates and not just those of the price equations because of the greater sensitivity of FIML estimates to model misspecification. To relate these findings to the literature on biases from measurement errors, the analysis applies OLS to (2) with both bilateral and multilateral price data. Inspection of the evidence reveals that reliance on multilateral price data underestimates the own-price effects and lowers their precision (lines 7 and 10), as predicted by Orcutt more than forty years ago.

5. Conclusions

This paper explains U.S. import demand without relying on the most restrictive assumptions embodied in fifty years of empirical work in trade: exogeneity of prices, absence of cross-price effects, autonomy of elasticities, and independence between spending and domestic decisions. Although several studies have addressed some of these limitations, no study has eliminated them all. To accomplish this goal, the paper assembles and estimates a simultaneous model explaining purchases and prices for U.S. expenditures. Spending behaves according to the Rotterdam model which,

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[21] Because the data for trade volumes are generated by deflating the value of trade flows by the corresponding bilateral price, estimating the system with multilateral prices, p_i^*, involves recomputing the data for trade volumes and replacing p_i with p_i^* in all of the equations of the model.

by design, embodies all of the properties of utility maximization. Pricing behaves according to the pricing-to-market hypothesis and allows exporters to update their pricing policy in response to changes in price elasticities. Parameter estimation relies on FIML and uses bilateral price data for 1965-1987.

By and large, the evidence speaks for itself. Ignoring the implications of optimization, simultaneity, and measurement errors carries important implications for key questions such as the responsiveness of U.S. imports to changes in income and prices and the extent to which volatility in nominal exchange rates has delinked U.S. trade from exchange rates. These findings, however, ignore intertemporal considerations and aggregation biases arising from the product composition of U.S. imports. These limitations are important but the process of eliminating them strengthens the point of the paper: that we should abandon the view that trade elasticities are autonomous parameters, that both cross-price effects and simultaneity biases are absent, and that expenditures on domestic and foreign goods can be studied independently of each other.

Appendix A: Chronology of Empirical Studies of U.S. Imports

This appendix lists papers reporting econometric estimates of trade elasticities for U.S. import demand. By design, the appendix excludes papers studying the structure of U.S. trade on the basis of factor-content (Bowen, Leamer, and Sveikauskas, 1987; Branson and Monoyios, 1977; Maskus 1983) or relying on non-parametric methods of estimation (Rousslang and Parker, 1984). Furthermore, the appendix focuses on features relevant for this paper: trade disaggregation, functional form, estimation method, and the presence of cross-price effects. The appendix also reports whether the data set includes bilateral or multilateral prices, the frequency of observation, and the period covered in estimation.

Table A1
U.S. Import Demand:
Chronology of Selected Studies

Study	Trade Disaggregation	Functional Form	Estimator[1]	Cross-price Effects	Price Data	Frequency;[2] Sample Range
1. de Vegh (1941)	Countries	Linear and log-linear	OLS	No	No price Effects	A; 1919-38
2. Addler (1945)	Dutiable and Duty-free	Linear	OLS	No	Multilateral	A; 1922-37
3. Addler, Schlesinger, and Westerborg (1952)	Countries and Commodities	Linear and log-linear	OLS	No	Bilateral	A; 1923-37
4. Liu (1954)	Countries and Commodities	Linear	OLS	Yes	Bilateral	A; 1923-37
5. Ball and Marwah (1962)	Commodities	Linear	OLS	No	Multilateral	A; 1948-58
6. Kreinin (1967)	Countries and Commodities	Log-linear	OLS	No	Cross-Sec.	A; 1955-57
7. Branson (1968)	Commodities	Shares in GNP	OLS	No	Multilateral	Q; 1955-1966
8. Adams, Eguchi, and Meyer-zu-Schlochtern (1969)	None	Linear	OLS	No	Multilateral	A; 1955-65
9. Houthakker and Magee (1969)	Countries and Commodities	Log-linear	OLS	No	Multilateral	A; 1951-66
10. Gregory (1971)	None	CES	OLS	No	Multilateral	Q; 1948-68
11. Price and Thornblade (1972)	Countries and Commodities	Log-linear	OLS	Yes	Bilateral	Q; 1964-69
12. Kreinin (1973)	Commodities	Log-linear	OLS	No	Multilateral	Q; 1964-70
13. Taplin (1973)	Commodities	Log-linear	OLS	No	Multilateral	A; 1953-70
14. Joy and Stolen (1975)	None	Log-linear	OLS	No	Multilateral	Q; 1953-71
15. Khan and Ross (1975)	None	Log-linear	OLS	No	Multilateral	S; 1960-72
16. Beenstock and Minford (1976)	None	Log-linear	OLS	No	Multilateral	Q; 1955-71
17. Goldstein and Khan (1976)	None	Log-linear	FIML	No	Multilateral	Q; 1955-70
18. Hooper (1976)	Commodities	Log-linear	OLS	No	Multilateral	Q; 1956-75
19. Marwah (1976)	Countries	Allocation	OLS	No	Multilateral	A; 1960-69
20. Murray and Ginman (1976)	None	Log-linear	OLS	No	Multilateral	Q; 1961-68

Table A1 (continued)

Study	Trade Disaggregation	Functional Form	Estimator[1]	Cross-price Effects	Price Data	Frequency;[2] Sample Range
21. Khan and Ross (1977)	None	Log-linear	OLS	No	Multilateral	Q; 1960-72
22. Mutti (1977)	Commodities	Log-linear	OLS	Yes	Multilateral	A; 1958-72
23. Clements and Theil (1978)	Commodities	Rotterdam	OLS	Yes	Multilateral	A; 1971-75
24. Hooper (1978)	Commodities	Log-linear	OLS	No	Multilateral	Q; 1955-77
25. Lawrence (1978)	Commodities	Log-linear	OLS	No	Multilateral	S; 1962-77
26. Theil and Clements (1978)	Commodities	Rotterdam	ML	Yes	Multilateral	A; 1921-70
27. Stern, Baum, and Green (1979)	None	Log-linear	OLS	No	Multilateral	Q; 1973-76
28. Stone (1979)	Countries and Commodities	Log-linear	OLS/2SLS	No	Bilateral	S; 1963-72
29. Wilson and Takacs (1979)	Countries	Log-linear	OLS	No	Multilateral	Q; 1957-71
30. Leamer (1981)	Commodities	Log-linear	OLS	No	Multilateral	Q; 1947-78
31. Grossman (1982)	Commodities	Log-linear	2SLS	Yes	Bilateral	Q; 1968-78
32. Haynes and Stone (1983)	None	Log-linear	OLS/Spectral	No	Multilateral	Q; 1955-79
33. Warner and Kreinin (1983)	Commodities	Log-linear	OLS	No	Multilateral	Q; 1957-70
34. Thursby and Thursby (1984)	None	Linear and log-linear	OLS	Yes	Multilateral	Q; 1955-78
35. Aw and Roberts (1985)	Countries	Allocation Translog	I3SLS	Yes	Multilateral	A; 1960-80
36. Kohli (1985)	Countries	Translog	SUR	Yes	Multilateral	A; 1960-79
37. Italianer (1986)	Countries and Commodities	Allocation CRESH	NLS	Yes	Bilateral	A; 1963-80
38. Shiells, Stern, and Deardorff (1986)	Commodities	Log-linear	2SLS	Yes	Multilateral	A; 1962-78
39. Helkie and Hooper (1987)	Commodities	Log-linear	OLS	No	Multilateral	Q; 1969-84
40. Kohli and Morey (1988)	Countries	Allocation CES	NLS	Yes	Multilateral	Q; 1960-79

Table A1 (continued)

Study	Trade Disaggregation	Functional Form	Estimator[1]	Cross-price Effects	Price Data	Frequency;[2] Sample Range
41. Marquez and McNeilly (1988)	Countries and Commodities	Log-linear	2SLS	No	Bilateral	Q; 1974-84
42. Parikh (1988)	Countries	Almost Ideal	ML	Yes	Multilateral	A; 1965-80
43. Deyak, Sawyer, and Sprinkle (1989)	Commodities	Log-linear	OLS	No	Multilateral	Q; 1958-83
44. Gagnon (1989)	Countries	Log-linear	ML	No	Multilateral	Q; 1973-85
45. Shiells (1989)	Countries and Commodities	Log-linear	2SLS	Yes	Multilateral	A; 1963-78
46. Cushman (1990)	Countries	Log-linear	OLS	Yes	Multilateral	Q; 1974-83
47. Kohli (1990)	None	Translog	I3SLS	Yes	Multilateral	Q; 1948-87
48. Lawrence (1990)	Commodities	Log-linear	OLS	No	Multilateral	Q; 1976-90
49. Marquez (1990)	Countries	Log-linear	OLS	Yes	Multilateral	Q; 1973-85
50. Marquez (1991)	Countries	Log-linear	FIML	Yes	Multilateral	Q; 1973-84
51. Shiells (1991)	Commodities	Log-linear	2SLS	No	Multilateral	Q; 1978-88

[1] *CES*: Constant Elasticity of Substitution; *CRESH*: Constant Ratios of Elasticity of Substitution Homogenous-Homothetic; *GLS*: Generalized Least Squares; *FIML*: Full Information Maximum Likelihood; *I3SLS*:Iterated Three Stage Least Squares; *ML*: Maximum Likelihood; *NLS*: Nonlinear Least Squares; *OLS*: Ordinary Least Squares; *SUR*: Seemingly Unrelated Regressors; *2SLS*: Two Stage Least Squares.

[2] A: Annual; Q: Quarterly; S: Semi-annual.

Appendix B: Implications of Spending Decisions for Pricing Equations

Following Gagnon and Knetter (1990, p. 6), let

(B.1) $$\varphi_{it} = [\partial \ln\psi_{ii,t} / \partial \ln p_{it}] \,/\, [\psi_{ii,t} - 1 + \partial \ln\psi_{ii,t} / \partial \ln p_{it}],$$

where φ_{it} is the same as Gagnon and Knetter's β_i and $\psi_{ii,t}$ is the uncompensated price elasticity in absolute terms. For the Rotterdam model, this elasticity is

(B.2) $$\psi_{ii,t} = -\pi_{ii} / w_{it} + \mu_i .$$

Taking logs and differentiating (B.2) yields

(B.3)
$$\begin{aligned}
\partial \ln\psi_{ii,t} / \partial \ln p_{it} &= [\pi_{ii} / (w_{i\,t}^{2}\psi_{ii,t})]\,[\partial(w_{it})/\partial \ln p_{it}] \\
&= [\pi_{ii} / (w_{it}\psi_{ii,t})]\,[\partial \ln(w_{it})/\partial \ln p_{it}] \\
&= [\pi_{ii} / (w_{it}\psi_{ii,t})]\,[1 + \partial \ln(q_{it})/\partial \ln p_{it}] \quad \text{[because } w_{it} = p_{it}q_{it} / y_t\text{]} \\
&= [\pi_{ii} / (w_{it}\psi_{ii,t})]\,[1 + \pi_{ii} / w_{it} - \mu_i] \quad \text{[use equation 2]} \\
&= \pi_{ii}(1 - \psi_{ii,t}) / (w_{it}\psi_{ii,t}) \ .
\end{aligned}$$

Substituting (B.3) into (B.1) gives

(B.4)
$$\begin{aligned}
\varphi_{it} &= [\pi_{ii}(1 - \psi_{ii,t}) / (w_{it}\psi_{ii,t})] / \{\psi_{ii,t} - 1 + [\pi_{ii}(1 - \psi_{ii,t}) / (w_{it}\psi_{ii,t})]\} \\
&= [\pi_{ii}(1 - \psi_{ii,t}) / (w_{it}\psi_{ii,t})] / \{[1 - \psi_{ii,t}][(\pi_{ii} / (w_{it}\psi_{ii,t})) - 1]\} \\
&= \pi_{ii} / [\pi_{ii} - w_{it}\psi_{ii,t}] \\
&= \pi_{ii} / [\pi_{ii} - w_{it}(-\pi_{ii} / w_{it} + \mu_i)] \\
&= \pi_{ii} / [2\pi_{ii} - \mu_i w_{it}] \ .
\end{aligned}$$

Reconciling differences in the definition of the exchange rate between Gagnon and Knetter's paper and this one involves setting $\delta_{it} = \varphi_{it} - 1$. Thus

(B.5) $$\delta_{it} = \{ [\pi_{ii} / (2\pi_{ii} - \mu_i w_{it})] - 1\},$$

which is equation (9) in the text. If the utility function exhibits preference independence (equation 7), then $\pi_{ii} = \phi\mu_i(1 - \mu_i)$. Substituting this expression into (B.5) results in

(B.6)
$$\begin{aligned}
\delta_{it} &= \{[\phi\mu_i(1 - \mu_i)/(2\phi\mu_i(1 - \mu_i) - \mu_i w_{it})] - 1\} \\
&= \{[\phi(1 - \mu_i)/(2\phi(1 - \mu_i) - w_{it})] - 1 \},
\end{aligned}$$

which is the passthrough embodying both Preference Independence and profit maximization.

Appendix C: Data Sources

This appendix describes the construction of the data and identifies the associated sources. Variables are classified as belonging to one of three categories: International Trade and Finance, Domestic Production and Prices, and Factor Prices.

C.1 International Trade and Finance

Imports: Values and Prices. Data for both the value and the price of imports come from the Commission of the European Communities and were provided by Alexander Italianer. These data are constructed from statistics collected by the OECD for trade values and trade volumes disaggregated at the five-digit level for a total of 1924 commodity categories for each trading partner. The value of these trade flows is measured in U.S. dollars on a CIF basis. For each of these categories, Italianer constructs a bilateral unit value measured in dollars; these unit values are then aggregated using Fisher Ideal Indexes into several commodity groupings including bilateral trade aggregated across commodities which is the one used in this paper. By relying on bilateral prices, the analysis recognizes that exporters might exploit elasticity differentials across export markets and result in different prices for the same product. Given the commodity-coverage involved in the construction of the data, the compilation of the associated prices is slow which means that they are available at an annual frequency and released with long delays. The data used in this study cover 1965-1987. For the purposes of this paper, imports from both OPEC and Centrally Planned Economies are excluded from the analysis.

Trade Volumes. International trade volumes (q_i, $i=0,\ldots, n-1$) are estimated as the value of trade flows deflated by the corresponding bilateral price and are expressed on a per-capita basis. The data for population come from the U.S. Department of Labor (Bureau of Labor Statistics, **Employment and Earnings**, table 1, column 1).

Tariffs. Data for the multilateral tariff rate are constructed as the ratio of Custom Duties (NIPA, table 3.2) to the total value of U.S. imports (custom basis) found in the U.S. Department of Commerce (**Survey of Current Business**, International Transactions, table 1).[22]

Exchange Rates. The bilateral dollar exchange-rates for Canada, Germany, and Japan come from the **International Finance Statistics**, Yearbook, 1989, published by the International Monetary Fund (line AF).

Multilateral Prices. Data for multilateral prices are constructed as the ratio between the nominal and the real value of exports of goods and services on a national income accounting basis. For Canada, the data come from the **Canadian Economic Observer**, Section 1, table 1.3; for Germany, the data source is the Supplement to the Monthly

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[22] NIPA stands for National Income and Product Accounts and the data appear in the *Survey of Current Business* published by the U.S. Department of Commerce.

Report of the Bundesbank, Section 4, table 1. For Japan, the data come from the **Economic Statistics Monthly** published by the Bank of Japan (127(3)<94161).

C.2 Domestic Production and Prices

The value of domestic production consumed domestically is measured as the nominal value of GNP minus the nominal value of exports of goods and services. Both the nominal and real value of exports of goods and services are measured on a national income accounting basis (NIPA, tables 1.1 and 1.2). Domestic purchases of domestic goods in real terms, q_n, are measured as the real value of GNP minus the real value of exports of goods and services on a per-capita basis (NIPA, table 1.2). The price for the domestic product (p_n) is estimated as the ratio between the nominal and real values of domestic spending on domestic goods.

C.3 Factor Prices

Labor costs in foreign countries, in foreign currency, are unit labor costs in manufacturing; the data come from the U.S. Department of Labor (Bureau of Labor Statistics USDL-90383, table 2, p.21). The wage rate for the United States is constructed as the ratio between Compensation of Employees (NIPA, table 1.14) and employment from the U.S. Department of Labor (Bureau of Labor Statistics, **Employment and Earnings**, table 1, column 6). The rental rate of capital is measured as

$$r_{nt} = \prod_{j=0}^{t} pk_j \Big/ \prod_{j=0}^{1982} pk_j ,$$

where $pk_t = (1-\tau_{pt} - \tau_{ct})[\, i_t - \delta_t - \pi_{kt}\,]$, τ_{pt} is the average personal tax rate constructed as the ratio between personal income tax receipts (NIPA, table 3.2) and nominal GNP (NIPA, table 1.1); τ_{ct} is the average corporate tax rate measured as the ratio between corporate profit tax accruals (NIPA, table 3.2) and nominal GNP (NIPA, table 1.1); i_t is the market yield on U.S. Treasury securities with 10 year maturity (Board of Governors of the Federal Reserve System, statistical release H15); δ_t is the depreciation rate of the non-residential capital stock; and π_{kt} is the inflation rate of the deflator for non-residential investment (NIPA, tables 1.1 and 1.2).

Appendix D: Computation of Standard Errors

D.1 Standard Errors for the Slutsky Coefficients. Assuming Preference Independence means that the Slutsky coefficients are nonlinear functions of the marginal budget shares and ϕ:

$$\pi_{ij} = -\phi\, \mu_i \mu_j \qquad (7)$$
$$\pi_{ii} = \phi\, \mu_i (1-\mu_i)$$

$\forall\, i,j \ni i \neq j$. This dependence means that the distribution of $\hat{\pi} \equiv (\hat{\pi}_{11} \ \ldots\ \hat{\pi}_{44})'$ is a nonlinear function of the distribution of $\hat{\Phi} \equiv (\hat{\mu}_1 \ \ldots\ \hat{\mu}_4 \ \hat{\theta}_1 \ \hat{\gamma}_1 \ldots \hat{\theta}_4 \ \hat{\gamma}_4 \ \hat{\phi})'$. Thus the estimation of the variance-covariance matrix of $\hat{\pi}$ involves complicated numerical integration. To avoid these complications, the paper uses Monte-Carlo simulations as a stochastic analogue to numerical integration. Specifically, assuming that $\hat{\Phi} \sim N(\Phi,\ \Omega)$, the paper generates a random sample of $\hat{\Phi}$ as $\Phi^k = \hat{\Phi} + \hat{\Gamma}\, \xi^k$, where Φ^k is the kth drawing, $\hat{\Gamma}$ is the Cholesky decomposition of $\hat{\Omega}$ ($\hat{\Gamma}\hat{\Gamma}' = \hat{\Omega}$), $\hat{\Omega}$ is the sample estimate of Ω, and ξ^k is a vector of drawings from an independent standard normal distribution. For each of these drawings, the paper uses (7) above to compute the associated vector of the kth drawing of Slutsky coefficients $\pi^k \equiv (\pi^k_{11} \ldots \pi^k_{44})'$, where $\pi^k_{ij} = -\phi^k\, \mu^k_i\, \mu^k_j$ and $\pi^k_{ii} = \phi^k\, \mu^k_i\, (1-\mu^k_i)$. Repeating this procedure 500 times generates a random sample of π^ks: $\Pi = (\pi^1 \ldots \pi^{500})$, where Π is a 10x500 matrix of drawings. The variance-covariance matrix of $\hat{\pi}$, $\hat{\Omega}_\pi$, is calculated as $(1/500)[\Pi - \hat{E}(\Pi)][\Pi - \hat{E}(\Pi)]'$.

D.2 Standard Errors for the Eigenvalues of the Slutsky Matrix. The standard errors of the eigenvalues are calculated using a Taylor expansion:

$$(1/2)\{[\partial\hat{\lambda}/\partial\hat{\pi}]_{\mathrm{i}} + [\partial\hat{\lambda}/\partial\hat{\pi}]_{\mathrm{d}}\}\ \hat{\Omega}_\pi\ \{[\partial\hat{\lambda}/\partial\hat{\pi}]_{\mathrm{i}} + [\partial\hat{\lambda}/\partial\hat{\pi}]_{\mathrm{d}}\}'(1/2),$$

where $\hat{\lambda}$ is the 4x1 vector of estimated eigenvalues, $\hat{\pi}$ is the 10x1 vector of parameter estimates, $\hat{\Omega}_\pi$ is the estimated 10x10 variance-covariance matrix of $\hat{\pi}$, and $[\partial\hat{\lambda}/\partial\hat{\pi}]$ is the 4x10 matrix of partial derivatives. These derivatives are computed numerically by both increasing and decreasing each entry of $\hat{\pi}$. A subscript "i" means that the partial derivative is calculated by increasing the entries of $\hat{\pi}$; a subscript "d" means that the partial derivative is calculated by decreasing the entries of $\hat{\pi}$.

D.3 Standard Errors for Passthrough Coefficients. The passthrough coefficients reported in figure 2 are constructed as

$$\hat{\delta}_{it} = \{[\hat{\phi}(1 - \hat{\mu}_i)/(2\hat{\phi}(1 - \hat{\mu}_i) - w_{it})] - 1\},$$

which is a nonlinear and time-dependent function of the structural parameter estimates, two features that complicate the derivation of the associated standard errors. To address the resulting complications, the analysis implements the Monte-Carlo procedure outlined in ***D.1*** and generates a random sample of passthrough coefficients for each period as

$$\hat{\delta}^k_{is} = \{[\hat{\phi}^k(1 - \hat{\mu}^k_i)/(2\hat{\phi}^k(1 - \hat{\mu}^k_i) - w_{is})] - 1\}$$

where k=1,..., 500, s=1965-1987, and both $\hat{\phi}^k$ and $\hat{\mu}^k_i$ are the random drawings of $\hat{\phi}$ and $\hat{\mu}_i$ generated in section ***D.1***; these drawings are unchanged across time. Given the random sample $\{\hat{\delta}^k_{is}\}^{500}_{k=1}$, the analysis computes the mean and the scaled median absolute deviation as measures of the first and second moments, respectively, for each period s.

Appendix E: The Almost Ideal Demand System

E.1 Model Specification

Assuming that $u(\cdot)$ belongs to the class of PIGLOG preferences (see Deaton and Muellbauer, 1980b), the demand equations associated with the Almost Ideal Demand system are

(E.1) $$dw_{it} = \rho_i \, dln(y/P)_t + \sum_{j=0}^{n} \zeta_{ij} \, dlnp_{jt} + r_{it} \ , i=0,..., n,$$

where $$dP_t = \sum_{j=0}^{n} w_{jt} \, dlnp_{jt} \, ,$$

and r_{it} is a random disturbance.[23] The parameters ρ_i and ζ_{ij} measure the response of the ith expenditure share when real income and the ith price increase by one percent, respectively. A positive ρ_i means that individuals treat the ith good as a luxury; otherwise the ith commodity is a necessity. Note that the sign of ζ_{ii} is not restricted a priori and thus $\zeta_{ii}>0$ does not imply an upward sloping conditional demand curve. The income and compensated-price elasticities implied by (E.1) are $[1 + (\rho_i / w_{it})]$ and $[-1 + w_{it} + (\zeta_{ij} / w_{it})]$, respectively (see Green and Alston, 1990).

Although equations (2) and (E.1) embody the properties associated with utility maximization and do not treat elasticities as autonomous parameters, they are different in two important respects. First, as noted by Deaton and Muellbauer (1980b), the Rotterdam model is an approximation to the solution of the first order conditions for optimizing an arbitrary utility function whereas the Almost Ideal Demand system is the exact solution to a specific utility function. Second, in the absence of income effects, the Rotterdam model predicts that the income elasticity is zero whereas the Almost Ideal Demand system implies a unitary income elasticity. Similarly, in the absence of price effects, the Rotterdam model predicts a compensated price elasticity of zero whereas the Almost Ideal Demand system implies a compensated price elasticity equal to $(-1 + w_{it})$. These differences in predictions have potentially important implications for understanding U.S. imports and motivate this paper's interest in considering the Almost Ideal Demand system.

For the parameters of (E.1) to be consistent with utility maximization they must satisfy

(E.2) *adding-up* $$\sum_{j=0}^{n} \rho_j = 0,$$

(E.3) *homogeneity* $$\sum_{j=0}^{n} \zeta_{ij} = 0 \text{ for } i=0,..., n \, ,$$

---

[23] Equation (E.1) corresponds to equation (17) of Deaton and Muellbauer (1980a, p.317). The expression for dP_t in (E.1) uses *Stone's price index* whereas the expression implied by the Almost Ideal Demand system is

$$dP_t = \sum_{j=0}^{n} \alpha_j \, dlnp_{jt} + (1/2) \sum_{j=0}^{n} \sum_{k=0}^{n} \zeta_{kj} \, d[lnp_{kt} \, lnp_{jt}].$$

Deaton and Muellbauer (1980a, p. 320) find that the results using Stone's price index are sufficiently close to those using the above expression.

(E.4) *symmetry* $\zeta_{ij} = \zeta_{ji} \; \forall i,j \ni i \neq j$, and

(E.5) *quasiconcavity* $[\kappa_{ijt}]$ is negative semidefinite with rank n, $\forall t$,

where $\kappa_{ijt} = \zeta_{ij} + \rho_i \rho_j \; ln(y/P)_t - \delta_{ij} w_{it} + w_{it} w_{jt}$ and δ_{ij} is the Kronecker delta.

Note that, unlike the Rotterdam model, the quasiconcavity conditions of the Almost Ideal system, equation (E.5), depend nonlinearly on income and prices and, therefore, these conditions have to be tested for every period.

For parameter estimation, the analysis retains the specification of prices for U.S. made products but, given the change in the structure of preferences (Rotterdam versus Almost Ideal), the analysis modifies the price equation for U.S. imports. This modification involves applying the reasoning of appendix B while recognizing that the uncompensated own-price elasticity of the Almost Ideal Demand System is $[-1 - \rho_i + (\zeta_{ii} / w_{it})]$ (see Green and Alston, 1990). The corresponding price equation is

$$lnp_{xit} = \theta_i + \beta_i \, lnC_{it} + \{[\zeta_{ii} / (2\zeta_{ii} - (1+\rho_i) w_{it})] - 1\} lnE_{it} + \gamma_i \, lnp_{nt} + e_{it},$$

where e_{it} is a disturbance. Again, as in the Rotterdam model, the effect of exchange rate changes on import prices depends on the structure of preferences. Specifically, if price effects are absent (ζ_{ii}=0) then $\delta_{it} = -1$ which means that foreign producers stabilize prices in their currencies. But the absence of price effects in (E.1) means that the uncompensated price elasticity is $(-1 - \rho_{it})$ whereas the absence of price effects in the Rotterdam model means that the uncompensated price elasticity is $-\mu_i$. These differences suggest that explaining export prices without reference to the structure of preferences entails a loss of information.

E.2 Empirical Results

Based on annual data for 1965-1987, table E.1 presents the FIML estimates of the parameters of the spending and pricing equations associated with the Almost Ideal Demand System. Inspection of the results reveals several findings of interest. First, the estimated income effects are positive and significant for U.S. imports but negative and significant for U.S. made products. This finding suggests that U.S. imports behave as luxuries whereas U.S. products behave as necessities. This evidence is consistent with the results from the Rotterdam model where the estimated income elasticities for U.S. imports are greater than one (see section 4.1). Second, the pricing-to-market coefficient varies from 0.42 for U.S. imports from Canada to 2.12 for U.S. imports from Japan. This dispersion of pricing-to-market coefficients is much wider than the range of estimates found for the Rotterdam model (table 3). Third, with the exception of Germany, foreign labor costs exert a large and significant influence on the prices for imports from Canada and Japan. Overall, the estimates of the price equations differ markedly from the estimates based on the Rotterdam model. This sensitivity emphasizes the importance of specifying explicitly the structure of preferences when modeling import-price behavior.

Given these parameter estimates, figures E.1 and E.2 report the 95 percent confidence intervals for income and price elasticities computed as $[1 + (\hat{\rho}_i / w_{it})]$ and $[-1 + w_{it} + (\hat{\zeta}_{ij} / w_{it})]$, respectively. Inspection of the results reveals a tendency for income elasticities to decline over time (figure E.1), as does the evidence from the Rotterdam model. The results for the compensated price elasticities indicate that U.S. imports from Canada and Japan have price elasticities which tend to increase (in absolute terms). Moreover, the price elasticity for purchases of U.S. products is not significantly different from zero. These results are not consistent with the evidence from the Rotterdam model.

Table E.1
Coefficient Estimates for Spending and Pricing Equations: Almost Ideal Demand System FIML, 1965-1987*

	Spending Decisions					Pricing Decisions	
	Income Effects	Price Effects				U.S. Price	Labor Costs
		Canada	Japan	Germany	U.S.		
Canada	0.014356	0.006559	0.006725	-0.001131	-0.018940	0.422781	1.679770
	0.005700	0.000036	0.000275	0.000877	0.002804	0.475600	0.437174
Japan	0.027276		0.003053	0.000466	-0.010835	2.116880	5.161950
	0.008309		0.000034	0.000934	0.003163	0.298341	0.305878
Germany	0.010902			-0.000140	-0.001325	1.161910	-0.396891
	0.002662			0.000167	0.001637	0.212125	0.310390
U.S.	-0.103856				0.065882		0.879622
	0.018811				0.010967		0.024057

* For a given cell, the top entry is the FIML estimate and the bottom entry is the associated standard error. The estimation sample is 1965-1987. See appendix C for the computation of the standard errors.

Figure E.1

Income Elasticities of Almost Ideal Demand System

95% Confidence Intervals

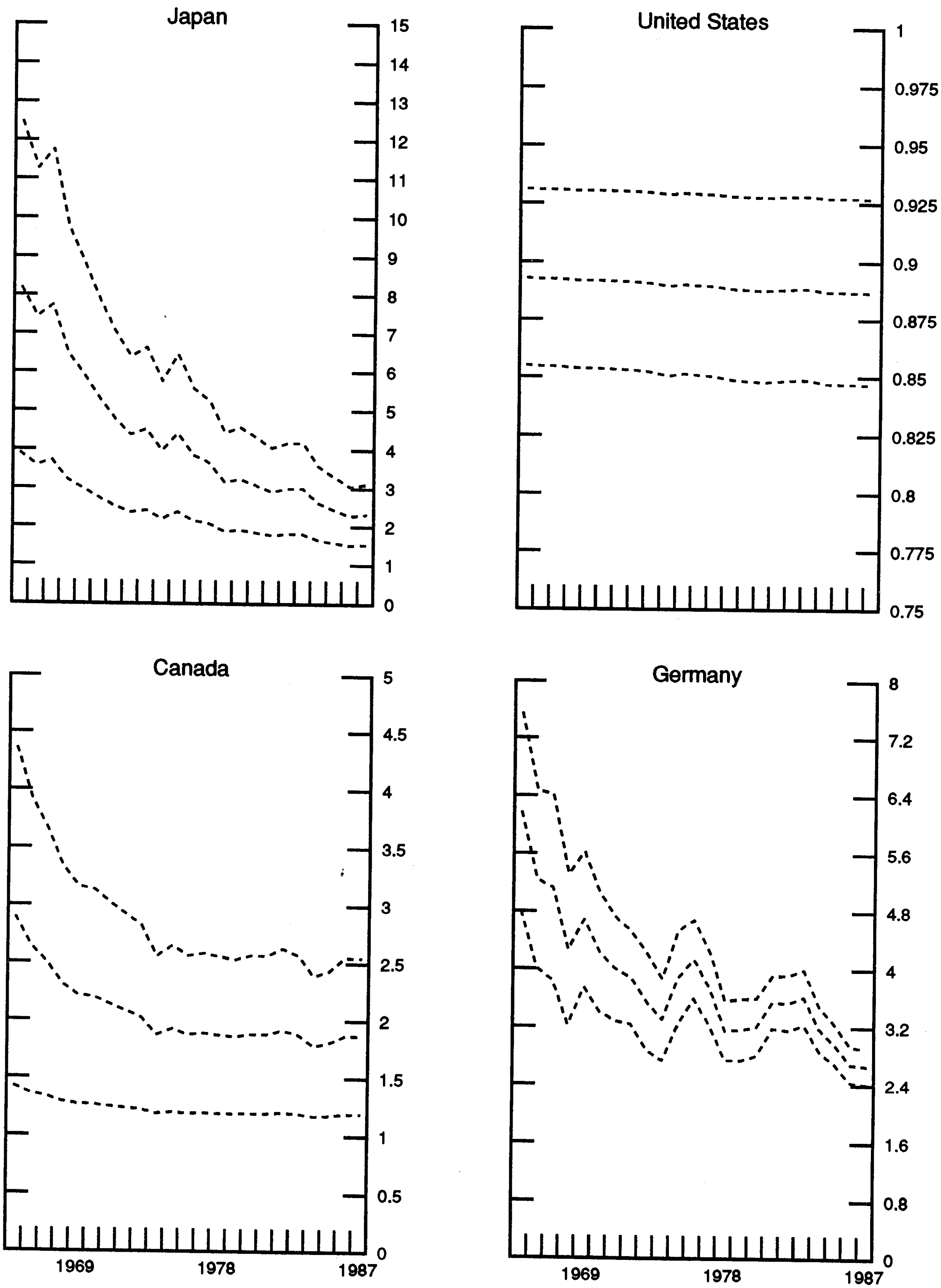
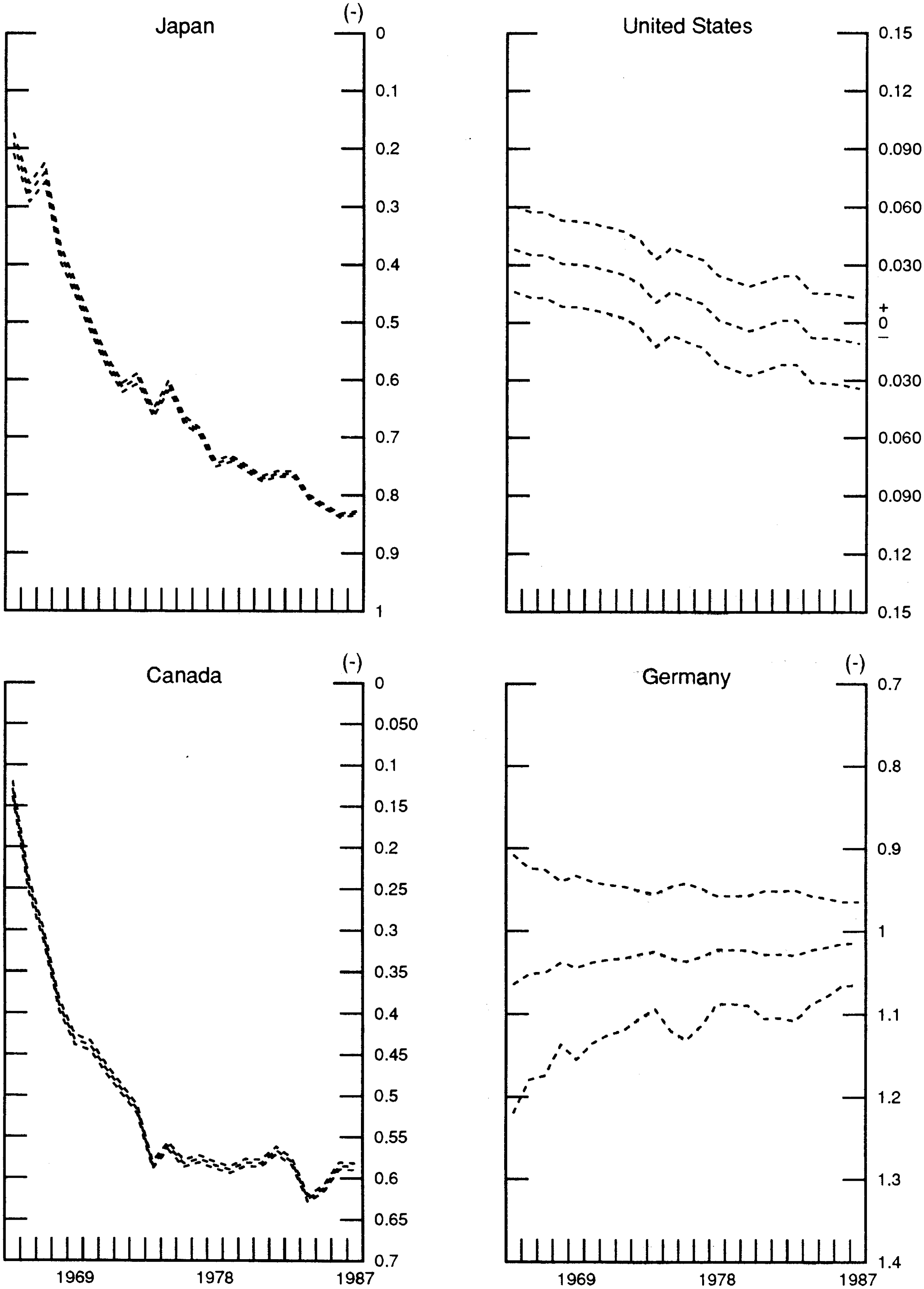


Figure E.2

Price Elasticities of Almost Ideal Demand System

95% Confidence Intervals

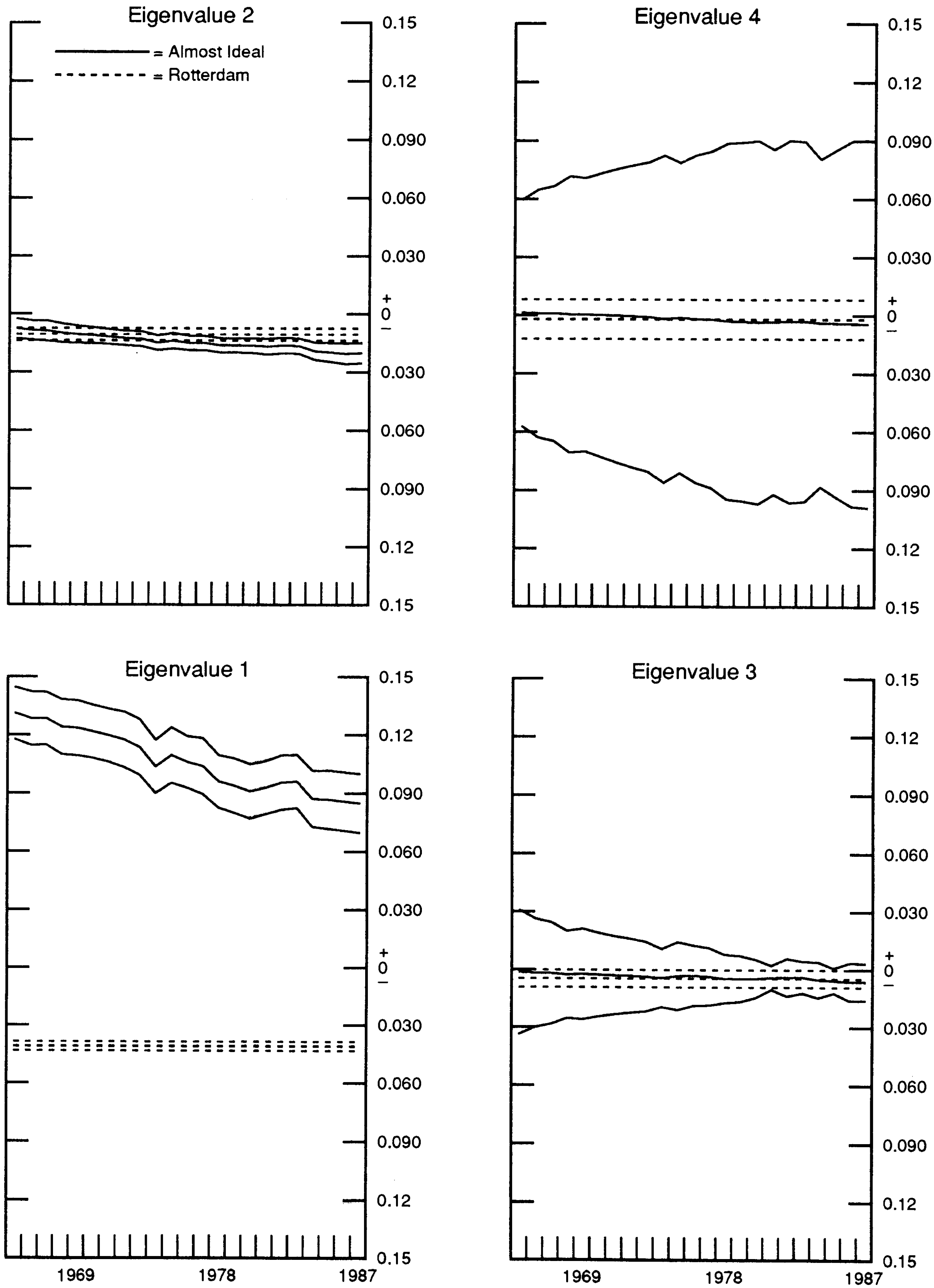


Differences in the estimated price elasticities between the Rotterdam model and the Almost Ideal Demand system raise the question of how to select between these two models. To this end, Deaton (1978) offers a convenient test to differentiate among non-nested demand systems. This paper does not implement Deaton's test because the data reject the negativity conditions (eq. E.5) as evidenced by one of the eigenvalues being positive for the entire sample (figure E.3); the data, however, cannot reject the restrictions implied by homogeneity and symmetry (eqs. E.3 and E.4). Note that the standard errors of the eigenvalues of (E.5) vary over time because the Slutsky coefficients are functions of income and prices. To compute these standard errors, I apply the formula shown in section D.2 (appendix D) for each observation. I want to emphasize that the rejection of these negativity requirements is conditional on the model chosen to explain prices. Alternative price formulations might reverse this result.

Figure E.3

Eigenvalues: Rotterdam and Almost Ideal Demand Systems

95% Confidence Intervals



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